Transmission of Sovereign Risk in the Euro Crisis*

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Abstract

We assess the role of financial linkages in the transmission of sovereign risk in the Euro Crisis. Building on the narrative approach by Romer and Romer (1989), we use financial news to identify structural shocks in a vector autoregressive model of daily sovereign CDS premia for eleven European countries. To estimate how these shocks transmit across borders, we use data on cross-country bank exposures to sovereign debt. Our results indicate that exposure to Greek sovereign debt and the debt of Greek banks constitute important transmission channels. All else being equal, the transmission rate to the country with the greatest exposure to Greece (1.22 percent of GDP) has been roughly 46 percent higher than the rate to the country with the least exposure (0.08 percent of GDP).

Keywords: Cross-country Transmission, Sovereign Risk, Financial Linkages, Euro Crisis, Narrative Approach

JEL Classification: F36, F42, F21, C30

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1 Introduction

In late 2009, with the global economy inching out of the Great Recession, the sovereign debt crisis hit Europe with a remarkable pace and vigor. Fears of sovereign insolvency initially developed in one peripheral country, Greece, but quickly spread to other European countries. In this paper, we assess the role of financial linkages in the transmission of sovereign risk in the Euro Crisis.

Financial linkages may spread sovereign risk across borders through different channels. Specifically, a looming Greek sovereign default may adversely affect banks in other countries according to their exposure to Greek sovereign debt. Through implicit guarantees, the troubled national banking systems constitute liabilities of the respective sovereign, thereby increasing sovereign credit risk. By the same token, cross-border interbank lending can matter: as the Greek sovereign debt crisis stresses the Greek banking system, counterparty risk rises for foreign creditors, affecting the financial health of the latter's countries.

Sketching this picture, Christine Lagarde, Managing Director of the International Monetary Fund, states that "[f]inancial exposures across the continent are transmitting weakness and spreading fear from market to market, country to country, periphery to core."[1] Likewise, concerns about risk propagation through financial connections have appeared regularly in the public debate.[2]

The current paper provides empirical evidence supporting these common beliefs. Specifically, we quantify the contribution of financial linkages to the transmission of sovereign risk, distinguishing in particular between transmission through exposure to public debt and through cross-border bank linkages. To this end, we identify financial shocks that originate in Greece and analyze how the sovereign default risk of European countries, measured by sovereign Credit Default Swap (CDS) premia, responds to these shocks. We then relate the responses to the cross-border exposure to Greek sovereign debt and debt of Greek banks.

Our results show that financial linkages significantly contribute to the transmission of sovereign default risk. In particular, bilateral exposure to sovereign debt and cross-border bank linkages constitute economically and statistically significant transmission channels.

[1] "Global Risks Are Rising, But There Is a Path to Recovery": Remarks at Jackson Hole, August 27, 2011

[2] See Financial Times: "Greek contagion fears spread to other EU banks", June 15, 2011.

Our results suggest that a 10-percent increase in the exposure to Greek debt increases the rate of cross-country transmission of sovereign risk by 3.94 percent. Similarly, a 10-percent increase in the exposure to debt of Greek banks implies that sovereign CDS premia react 1.47 percent more strongly to a Greek shock. A back-of-the-envelope calculation based on these numbers shows that, everything else being equal, the transmission rate to the country with the greatest exposure to Greece (1.22 percent of GDP) has been roughly 46 percent higher than the transmission rate to the country with the least exposure (0.08 percent of GDP).

Methodologically, we consider a vector autoregressive (VAR) model for country CDS premia, which includes global financial factors as exogenous variables. We adopt the narrative approach of Romer and Romer (1989) to identify exogenous shocks in one country, focusing specifically on information innovations from Greece or "Greek shocks". Within this framework, we estimate whether the exposure to Greek debt has contributed to the transmission of sovereign risk from Greece to other European countries. In doing so, we adopt a difference-in-difference approach and we test whether within-country changes in the response to Greek shocks correlate with within-country changes in exposure to Greek debt.

The narrative approach has the major advantage that identification is achieved without imposing a specific pattern for cross-country spillovers on the VAR model. This virtue is particularly important when analyzing cross-country transmission of risk because standard identification assumptions, e.g., short-run, long-run or sign restrictions, are difficult to defend in this context (see Rigobon (2002)). Furthermore, the narrative approach is especially suitable in the case of daily data, as the identification of the events becomes much cleaner and the problem of omitted variables - i.e., unobserved simultaneous shocks outside Greece - is sensibly reduced. Within this framework, we need to control for the heteroskedasticity of CDS premia between days of exceptional events in Greece and all other days. We accordingly construct a general least squares (GLS) estimator that allows us to derive efficient estimates of the effects at study.

Some concerns need to be addressed when pursuing the narrative approach. Importantly, CDS responses in other countries would be erroneously attributed to spillovers from Greece if simultaneous shocks to these countries or to global factors occurred on the days of

Greek news. We argue that our identification of Greek shocks and of risk spillovers through financial linkages are most likely unaffected by this problem for the following reasons.

First, we take special care in reviewing financial news to exclude from our analysis those Greek events that occurred simultaneously with relevant financial news in other countries.[3] Next, by controlling for global financial factors in the estimation, common shocks that affect sovereign risk through general market conditions are accounted for. Moreover our difference-in-difference approach is not likely to suffer from an omitted variable bias. Specifically, we would observe a bias in our estimates only if the change over time of the response to unobserved shocks correlates with the change in exposure to Greek debt. We claim that such a systematic link is rather implausible. Nevertheless, it could be that changes in exposure to Greek debt correlate with changes in countries general exposure to risk. In this case our results could be driven by the updating of investors' beliefs regarding systemic risk on days of Greek news rather than by the direct transmission of Greek shocks. To address this concern, we re-estimate our model, replacing exposure to Greek debt with exposure to total foreign debt. The results indicate that exposure to foreign debt from origins other than Greece is unrelated to the reaction of sovereign risk. We thus conclude that our estimates of risk spillovers through direct financial linkages with Greece are not likely to be driven by contemporaneous changes in global conditions or market sentiment.

Potential endogeneity problems are another source of concern. Clearly, the exposure to Greek debt reflects endogenous decisions of banks, which may well be related to sovereign risks. For our results to be spurious in our difference-in-difference approach, the increase in the sensitivity of a country's sovereign risk to Greek shocks should lead domestic banks to raise their exposure to Greek debt. We find this possibility hard to defend. Moreover, it is reasonable to assume that the change in exposure rates, which occurs quarterly in our data, is exogenous to changes in the cross-country responses of CDS premia on a limited number of days within each quarter.

The classic use of the narrative approach has also been disputed in the past. Romer and Romer's (1989) original identification strategy through episodes of monetary contraction has been criticized. The effects of the policy events selected using a narrative approach are said to be indistinguishable from those of omitted economic fundamentals (Hoover and

[3] A detailed description of the data and the selection procedure follows in Section 2.

Perez (1995)). Additionally, the events considered may constitute an endogenous policy response to previous movements in economic conditions (Leeper (1997); see also the reply by Romer and Romer (1997).[4] We argue that these standard lines of critique do not apply to our setup. First, the high frequency of our data relieves us of the specific problem of reverse causality: a daily blip in the CDS premium cannot reasonably be believed to cause major political action in Greece on the same day (compare Table A1 in the Appendix). Second, unlike Romer and Romer (1989), we measure the immediate effect of the selected Greek events, i.e., the change of CDS premia on the same day, instead of medium- and long-term effects that could be driven by macroeconomic fundamentals.[5]. Finally, we also argue that our identification is intact even when events are partially anticipated as we measure the rate of transmission based on the response of other European countries relative to the size of the contemporaneous shock in Greece. A partial anticipation of future shocks thus has no effect on the estimates of this relative transmission.

1.1 Literature Review

The mechanism of cross-country contagion and transmission of financial turmoil is at the core of a large empirical and theoretical literature. Calvo and Reinhart (1996) distinguish between three different concepts: first, response to common shocks; second, fundamental-based contagion, which operates via direct trade or financial linkages; third, pure contagion, which operates through portfolio rebalancing and investors' herding behavior.[6] Our paper addresses fundamental-based contagion, and is close to the spirit of Forbes (2004) and Forbes and Chinn (2004).[7] Contrary to the latter study, our results indicate that bank linkages constitute an important transmission channel.[8] Likely reasons for these differences

[4]Leeper (1997) notes that the events that Romer and Romer (1989) use to identify the response to monetary shocks (contractions of the FED) are predictable by past macroeconomic variables and that unpredictable changes do not generate responses that look like typical effects of monetary policy. Romer and Romer (1997) argue that the results in Leeper (1997) are due to overfitting.

[5]We also refer to the lag structure of the VAR here, which generally captures trends of the dependent data. See the description of our autocorrelation tests in Section 4.

[6]A necessarily incomplete list of papers includes Baig and Goldfajn (1999), Schinasi and Smith (2000), Kyle and Xiong (2001), Forbes and Rigobon (2002), Corsetti et al. (2005), Broner et al. (2006).

[7]See also Kaminsky and Reinhart (2000), Hernandez and Valdes (2001), Van Rijckeghem and Weder (2001, 2003).

[8]Analyzing stock market returns, Forbes and Chinn find that bank channels are significant only in some of their specifications.

are the relatively strong European financial integration and the use of time variation to identify spillovers traveling through cross-border bank exposures.

Inspecting the role of common shocks, Longstaff et al. (2011) show that global factors, such as the US stock market return or the VIX volatility index, account for a large fraction of the common variation in CDS premia around the world. Borri and Verdelhan (2011) have a model where adverse shocks to the financial center (e.g., the US) determine, through investors' risk aversion, a change of risk premia in the periphery. Furthermore, the change depends on the business cycles synchronization with the financial center. We see our results as complementary to these studies insofar as we control for global factors in our estimations.[9]

In terms of methodology, we are not aware of any paper in the contagion literature applying a narrative approach to the identification of country-specific shocks and their transmission. We acknowledge, however, that the narrative approach exhibits some commonalities with a classic event-study methodology, which has been previously applied to study the international transmission of financial crises (e.g., Forbes (2004)). This methodology usually considers a market model to explain the behavior of financial assets during normal times and then compute for specific event-window the abnormal returns, defined as deviations from the model predictions. In our study, we are forced to use a VAR model to account for the large autocorrelation of CDS premia during the Euro Crisis. Second, event studies are typically based on a single event-window of extended length rather than on a collection of daily events. Our definition of an event is thus more conservative since a clear identification of the exact origin of the event is much easier to motivate for short-lived single-day events. Third, we estimate the contribution of financial linkages in the spreading of abnormal returns by adopting a one-stage procedure rather than the two-stage procedure frequently adopted in event studies. Our procedure thus allows us to handle the heteroskedasticity of abnormal returns and to obtain efficient estimates.[10]

The unfolding Euro crisis is currently stimulating rich academic output, which can be covered here only partially. A large part of this literature has focused on the role of finan-

[9]Furthermore, in a VAR model, the response of risk premia to global factors can differ across countries, as suggested by Borri and Verdelhan (2011).

[10]Forbes (2004) alternatively uses a two-stage procedure, where the second-stage uses a weighted estimator, proposed by Sefcik and Thompson (1986), in order to handle efficiently the heteroskedasticity of abnormal returns.

cial linkages in the spread of the crisis. Bolton and Jeanne (2011) develop a comprehensive theoretical analysis of transmission of sovereign risk through an integrated banking system. The authors note that, although "diversification generates risk diversification benefits ex ante, it also generates contagion ex post." Using sovereign and bank CDS premia between 2007 and 2010 and a series of bank bailouts, Acharya et al. (2011) provide evidence that a weak banking sector increases the default risk of the sovereign, showing in particular "that the announcement of financial sector bailouts was associated with an immediate, unprecedented widening of sovereign CDS spreads and narrowing of bank CDS spreads."[11] In line with these findings, Dieckmann and Plank (2010) find "a private-to-public risk transfer through which market participants incorporate their expectations about financial industry bailouts and the potential burden of government intervention." Focusing on the role of financial news, Bhanot et al. (2011) report results that "point to the role of news announcements and the banking channel as important transmission channels in the crisis period." Finally, Arezki et al. (2011) find that "sovereign rating downgrades have statistically and economically significant spillover effects both across countries and financial markets." The authors discuss several channels of spillover of sovereign risk across countries, pointing especially to "the holding of foreign sovereign debt by domestic banks..." and to the claims of banks "on banks in other countries" (see also Blundell-Wignall and Slovik (2010) and Sy (2010)).

Our contribution to this empirical literature is twofold. First, building on the documented roles of bank-to-bank and bank-to-sovereign linkages as transmission channels of sovereign risk, we take the broader perspective of the country level and analyze cross-country spillovers and the role the national exposure rates in the transmission of sovereign risk. Second, we evaluate the transmission mechanism through the careful identification of financial shocks that originated in Greece and through the time variation in national exposure rates.

The remainder of this paper is structured as follows. Section 2 describes our data in detail, and Section 3 lays out the empirical framework. Section 4 presents and discusses our main findings, and Section 5 concludes.

[11] Balteanu and Erce (2011) study the feedback effects between bank crises and sovereign defaults in emerging economies.

2 Data

Our analysis requires a combination of three main data types: (i) a measure of sovereign default risk; (ii) a measure of bilateral financial linkages; and (iii) a classification of shocks that identifies, in particular, those of Greek origin. The key sources for the first and second types of data are Datastream and the BIS, respectively. We compile the third on our own.

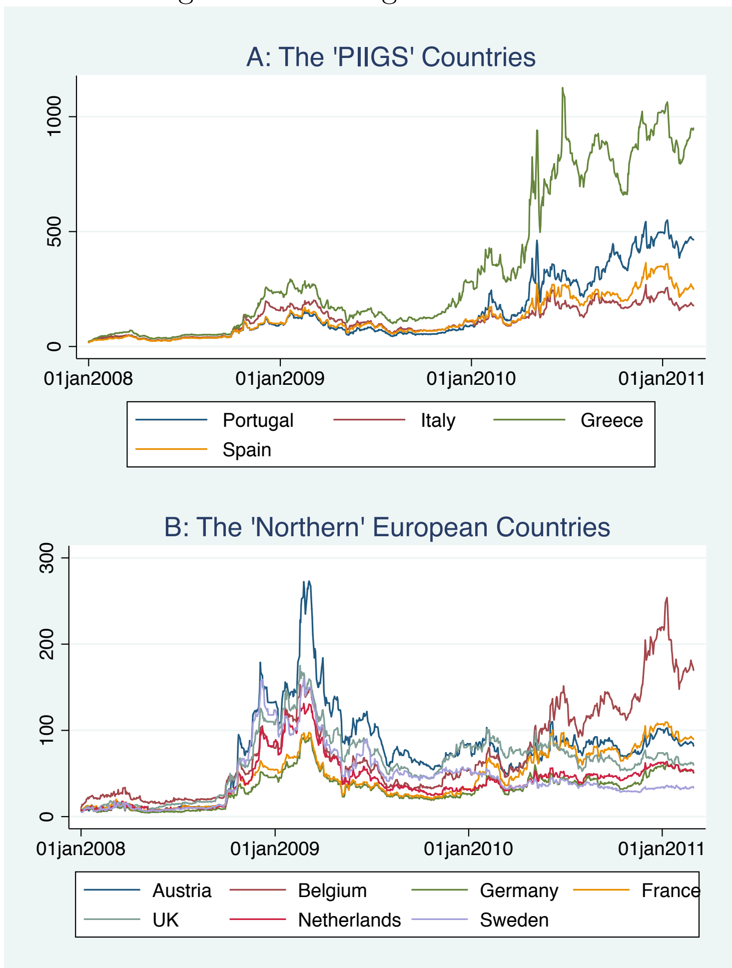
Sovereign default risk. Our measure of a country's default risk is a five-year sovereign Credit Default Swap (CDS) premium, collected by CMAN and provided by Bloomberg. A CDS is essentially an insurance contract between two counterparties, typically traded over the counter. It transfers credit risk from one party, the buyer of protection, who pays "a regular fixed premium to the seller of protection in return for compensation contingent on the occurrence of a specified credit event" (Barclays Capital 2010). A credit event, in turn, is a general form of partial or full default by the borrower. In the case of a sovereign CDS contract, the borrower is the sovereign of a country, and the underlying debt security is a government bond of a certain specified type and duration. The most frequently traded sovereign CDS contracts are typically five-year on-the-run government bonds, which have commonly been monitored, quoted and commented on in the course of the recent debt crisis.[12] Our analysis relies on this particular specification of sovereign CDS contracts.

Since 22 June 2009, the European CDS market has applied the Standard European Contract (STEC) that fixes trading conventions. These specify that: (i) quotations are for one of four fixed coupons: 25bp, 100bp, 500bp and 1000bp; (ii) the definition of a credit event comprises restructuring; and (iii) the premium leg incorporates a full first coupon, and accrued interest is paid at inception. Moreover, quoted spreads assume a flat credit curve to calculate the up-front payment and a recovery rate conventionally fixed at 40% for Western sovereigns (see Barclays Capital 2010). Although CDS premia do not capture default risk perfectly, recent literature has documented that they nevertheless constitute a quite reliable measure and are certainly among the best available (Pan and Singleton (2008), Stulz (2010) and Fontana and Scheicher (2010)).

Figure 1 plots daily five-year sovereign CDS premia separately for "PIGS" countries (Portugal, Italy, Greece and Spain) and eight other "Northern" European countries for

[12] See Packer and Suthiphongchai (2003), available at http://www.bis.org/publ/qtrpdf/r_qt0312g.pdf.

Figure 1: Sovereign CDS Premia



Five-year sovereign CDS premia of selected European countries at daily frequency from January 2008 to March 2011. Units of the vertical axis are basis points (a level of 100 bp implies that it costs 0.01 Euro per year to protect one Euro of outstanding debt over the five-year period). Source: Bloomberg.

the period from January 2008 to March 2011.[13] These graphs are familiar to the reader of financial news: the curves not only track financial troubles at the height of the Great Recession, but also illustrate the buildup of the Euro crisis; in particular, the market's assessment of a potential Greek sovereign default following the budgetary announcements

[13]Ireland, the additional I in the usual PIIGS acronym, is excluded from our sample for the following reason. In the last quarter of 2010 we observe a sharp decline in the exposure of Irish banks to Greek debt. This drop is likely to be caused by the creation in late 2010 of a nationalized "bad bank absorbing troubled assets owned by Irish banks, without actually changing the exposure to Greece. We are concerned that the inclusion of Ireland might bias our results.

by Greek authorities in fall 2009.

The figure suggests strong co-movements of the countries CDS premia. Indeed, Table 1 reports pairwise correlations of log-changes for the fourteen countries. These pairwise correlations range between 0.6011 (Sweden and Greece) and 0.8952 (Italy and Spain), with an average of 0.7290. In addition to the obvious effects of common factors (see Fontana and Scheicher (2010)), part of that correlation is likely to be generated by spillover effects from one troubled country – e.g., Greece – to other countries in the sample.

Financial linkages. Bilateral banking linkages are measured using data from the Consolidated Banking Statistics of the BIS. This dataset covers banks' claims to non-residents, broken down by nationality of borrower and creditor and four types of debt: (i) Public Debt; (ii) Private Bank Debt; (iii) Private Non-Bank Debt; and (iv) Other Debt (see BIS (2006)). Reporting entities are bank head offices, including the exposures of their foreign affiliates (i.e., subsidiaries and branches); reported data cover financial claims on the balance sheet, consolidated on a worldwide basis within each banking group (i.e., inter-office positions are netted out).[14] The final dataset provides a matrix of cross-country bilateral (gross) bank exposures, evaluated at market prices.[15]

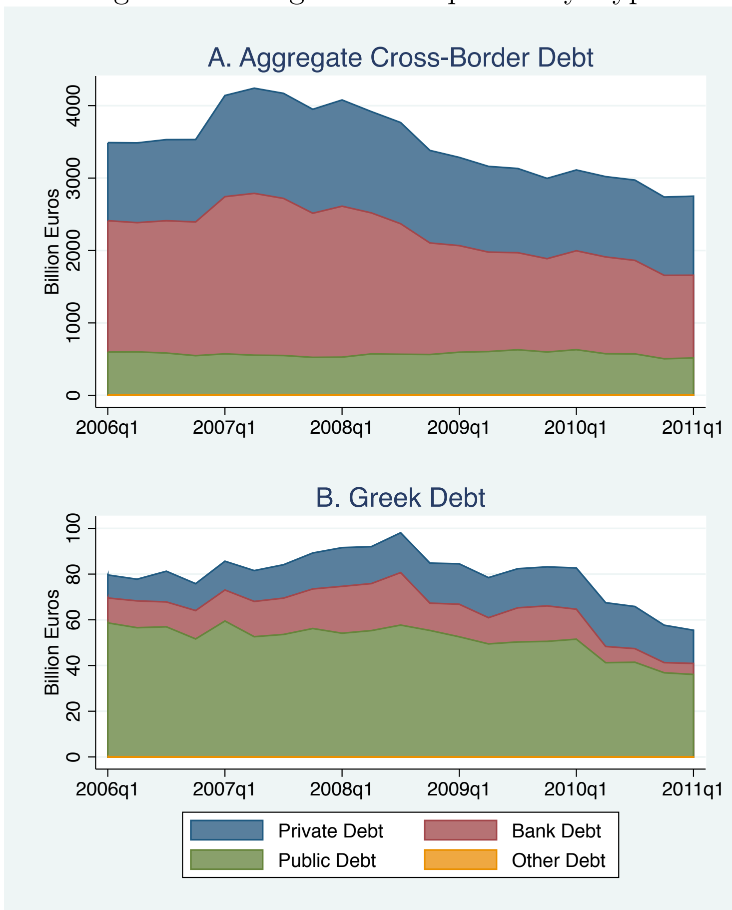
The BIS reports the data in US dollars after converting the positions originally reported in other currencies using end-of-quarter exchange rates. As most of the countries in our sample belong to the Euro Area, we convert the data back into Euros using the end-of-quarter exchange rate provided by the International Financial Statistics (IFS) of the IMF. This conversion allows us to remove potential valuation effects that are common to all countries, which might introduce an exogenous variation in our measure of financial linkages and thus affect our estimates. Finally, we normalize banks' exposure by the average GDP of the corresponding country to control for different country sizes. The data on GDP refer to the 2009-2010 period and come from the IFS.

Figure 2, top panel, illustrates gross cross-border exposure of BIS-reporting banks,

[14]BIS (2006) specifies that balance-sheet relevant instruments include "certificates of deposit (CDs), promissory notes and other negotiable paper issued by non-residents, banks' holdings of international notes and coins, foreign trade-related credits, claims under sale and repurchase agreements with non-residents, deposits and balances placed with banks, loans and advances to banks and non-banks, holdings of securities and participations including equity holdings in unconsolidated banks or non-bank subsidiaries."

[15]In addition, BIS (2006) reports a number of valuation practices recommended to reporting entities to ensure the cross-country comparability of data. In addition to the use of market prices to evaluate claims, these practices specify how to report arrears, provisions and the write-off of claims.

Figure 2: Foreign Debt Exposure by Type



Exposure of BIS reporting banks to European debt (top panel) and Greek debt (bottom panel) by type of debt: Public, Bank and Non-Bank Private, and Other. Quarterly data are reported on an immediate borrower basis (for a comparison between immediate and ultimate borrower bases, see the Appendix). The set of 11 countries includes: Austria, Belgium, France, Germany, Greece, Italy, Netherlands, Portugal, Spain, Sweden and UK. Source: BIS, Consolidated Banking Statistics.

aggregated across the eleven European countries in our sample, by type of debt. It shows that, during the height of the financial crisis, the total foreign positions were rewound, in particular those of bank debt. With the onset of the year 2010, the downward trend has continued, but moderately so. The picture is different when focusing on Greek foreign debt to the remaining ten countries (bottom panel). Here, the largest decrease in debt exposure occurred during the year 2010, with positions in sovereign debt dropping the most in absolute terms.[16]

[16]In principle, the decrease in the exposure to Greek bonds may result from valuation effects. Thus,

It is important to notice that neither the levels nor the trends in the exposure to Greek debt are uniform across all countries. Table 2a shows that within the set of ten European countries, average exposures vary between roughly 0.08 and 1.22 percent of GDP (for 2009 and 2010). Table 2b reports summary statistics of the evolution of exposure to Greek debt following the fourth quarter 2009, for which levels are normalized to 100.[17] Panel A reports numbers for total debt by quarter. The last column thus shows that, on average, the exposure to Greek debt was 67.5% of the initial level. The country for which exposure changed the least reduced its exposure by 0.2% (100-99.8), and the largest reduction amounted to 77.2%. Panel B summarizes exposures for Greek public debt. Up to 2010Q4, the average country reduced its exposure to 73.0% of the initial level. The maximum reduction was 82.3%, and one country actually increased its exposure by 13.1%. Finally, panel C reports the numbers based on Greek bank debt. Between 2009Q4 and 2010Q4, the average country reduced Its exposure to 32.0% of the initial level. The maximum reduction reached 97.7%, and at least one country increased its exposure (by 4.2%). Our aim is to exploit these cross-country differences in time variation to extract information about how sovereign risk is transmitted via exposure to the different classes of Greek debt.

These data are reported on the immediate borrower basis and do not include credit derivatives and guarantees, which could shift risk exposure across countries. Thus, the accuracy of effective exposure is a potential source of concern. Unfortunately, the availability of data on the ultimate risk basis is very limited. Not only would the use of the ultimate borrower basis limit our sample, but Greece would drop out of it. A comparison between the data on the immediate and ultimate basis shows that the two measures covary very closely.

Greek Shocks. The third type of required information is a list of those days when financially relevant news was dominated by information from Greece. For our identification strategy, it is essential that we are able to claim that Greece is the only origin of a first-order shock. Pioneering the "“narrative approach, ,” Romer and Romer (1989) stress that

to ensure international comparability, the BIS recommends "that banks' international claims be valued at market prices" but acknowledges that, where "market values are not appropriate, contractual or nominal values should be used." Moreover, differences across countries arise because "loans in the banking books, which in principle should be assigned nominal value, should be valued in accordance to the reporting countries' accounting standards." (BIS 2006)

[17]The beginning of the Greek sovereign debt crisis can be dated to 2009Q4.

potentially severe problems may arise when isolating the shocks due to the judgmental and retrospective nature of the selection process. To reduce the unconscious bias in this selection process, we keep our selection as mechanical as possible. Specifically, we first construct a timeline describing the Euro crisis by merging three sources: the Financial Times (Interactive Timeline: Greek Debt Crisis), the Wall Street Journal (Europe's Debt Crisis - Timeline), and Reuters (Europe's Debt Crisis Timelines).[18] The joint timeline compiled from this information covers a period starting November 5, 2009, and ending November 28, 2010, and includes those days that appear to be relevant to the authors of at least one of these timelines. We mechanically extract the days that contain news from Greece. We then limit the resulting list of days by excluding all days when financially relevant news items are reported from countries other than Greece. In addition, we perform a second round of elimination by searching for potential overlapping shocks originating in other countries based on the Lexis-Nexis database. The events thus collected are then classified as days of positive shocks or negative shocks, depending on whether we expect the Greek CDS premium to either increase or decrease on that day. The final list of events comprises 14 information shocks identified as Greek shocks and is reported in Table A1 along with a short description of the corresponding event.

Initial inspection of the data reveals two important aspects of CDS premia on the days classified as Greek events. First, on these days, we observe large movements in the Greek CDS premium relative to the rest of the sample. Table 3a reports summary statistics of daily log-changes in the Greek CDS premium, partitioning the sample into days of positive Greek shocks, days of negative Greek shocks and all remaining days, respectively. On average, the Greek CDS premium varies by more than 7% during Greek events, whereas the average change in the rest of the sample is 0.26%. Greek events have also coincided with episodes of extreme variation in the Greek CDS premium: the largest increase in the Greek CDS premium (27.00%) occurred on a day of positive Greek shocks. Similarly, the second largest drop in the Greek CDS premium (-15.70%) is associated with a Greek event.[19]

[18]See http://www.ft.com/intl/cms/s/0/003cbb92-4e2d-11df-b48d-00144feab49a.html#axzz1eQwHrlj, http://online.wsj.com/article/SB10001424052748704448304575195863350731920.html and http://www.reuters.com/article/2010/08/25/eurozone-crisis-events-idUSLDE67O0YD20100825

[19]Notice that the largest drop in the Greek CDS premium (-47.10%) coincides with the joint EU-IMF announcement of a 110bn Euro bailout package for Greece. However, we had to exclude this announcement from the list of pure Greek events as the EU Council agreed on the same day to establish the first European

Looking at the other European countries, we find that exposure to Greek debt above the country median is on average associated with larger CDS variation during Greek events. Table 3b reports summary statistics of the log-changes in the CDS premia of the 10 European countries in our sample by rate of exposure to Greek debt (above or below the country median) and by type of Greek event (positive or negative). As shown in Panel A, total exposure above the country median is associated on average with a 3.03% increase in CDS premia on the days of positive Greek shocks and a 2.99% decrease on the days of negative Greek shocks. These numbers shrink to 1.85% and 0.49%, respectively, when exposure is below the median. Panels B and C show similar patterns in the cases of exposure to public debt and to private bank debt, respectively. Overall, these numbers suggest that the within-country variation in exposure to Greece could affect the transmission of Greek shocks to other European countries. In the next section, we describe the empirical framework that we adopt to test this hypothesis.

3 Empirical Framework

In this section, we discuss our approach to identifying sovereign risk spillovers across countries. We start by considering the n-dimensional vector autoregressive model

$$y_t = \Phi_y(L)y_{t-1} + \Phi_x x_t + u_t$$

where y_t is a vector of CDS premia for each country, $\Phi_y(L) = \sum_{j=1}^{J} \Phi_{y,j} L^{j-1}$ is a polynomial of $n \times n$ matrices in the lag operator, x_t is a m-dimensional vector of exogenous variables that includes the constant, and u_t is a vector of innovations. In a more compact form, this model can be written as

$$y_t = \Phi z_t + u_t, \tag{1}$$

where $\Phi = [\Phi_{y,1}, \ldots, \Phi_{y,J}, \Phi_x]$ is a $n \times p$ matrix and $z_t = [y'_{t-1}, \ldots, y'_{t-J}, x'_t]'$ is a $p \times 1$ vector, where $p = nJ + m$.

The innovations u_t are jointly correlated reflecting cross-country spillovers. Following standard practice, we assume

$$u_t = B\epsilon_t, \tag{2}$$

stabilization mechanism with the aim of safeguarding financial stability in all member states.

where ϵ_t is an i.i.d. vector of country-specific shocks with zero mean. The matrix B instead reflects the rate of transmission of these shocks across countries. Without loss of generality, we normalize the diagonal elements of B to one.

3.1 Identifying Spillovers using a Narrative Approach

As part of our identification strategy, we decompose each country-specific shock into a signed mean shift, which shall reflects "exceptional" events, and a residual noise, which shall reflect "normal" shocks. Specifically, for each component i, we assume

$$\epsilon_{i,t} = d_{i,t}\xi_i + \nu_{i,t}$$

where $d_{i,t} \in \{-1, 0, 1\}$ is a signed indicator function which is independent across countries; ξ_i is a time-invariant parameter; and $\nu_{i,t}$ is independent of $d_{i,t}$ and i.i.d. across countries and time.[20] Applying the same decomposition to each country, we can write

$$\epsilon_t = D_t\xi + \nu_t$$

where D_t is a diagonal matrix with entries $d_{i,t}$. By adopting the "narrative approach" of Romer and Romer (1989), we are able to identify the dates on which an exceptional event occurred in Greece but in no other country. Using the signed indicator $\mathbf{1}_{gr,t}$ to denote the days of Greek events, we can thus write

$$\epsilon_t = \mathbf{1}_{gr,t}\xi^{gr} + (1 - |\mathbf{1}_{gr,t}|)\ \xi_t^{ngr} + \nu_t$$

where ξ^{gr} denotes the vector of exceptional shocks in Greek events and thus is zero except for its Greek component ξ_{gr}. Meanwhile, $\xi_t^{ngr} \equiv D_t\,\xi$ reflects the random occurrence of exceptional shocks in other countries in all other periods.

Combining (**??**) and (2), we can rewrite our original model (1) as

$$y_t = \Phi z_t + \beta \cdot \mathbf{1}_{gr,t} + \tilde{u}_t, \tag{3}$$

where $\beta \equiv B\xi^{gr}$ and $\tilde{u}_t$ is a modified residual defined as

$$\begin{aligned}
\tilde{u}_t &\equiv \mathbf{1}_{gr,t}u_t^{gr} + (1 - |\mathbf{1}_{gr,t}|)u_t^{ngr} \\
u_t^{gr} &\equiv B\nu_t \\
u_t^{ngr} &\equiv B(\xi_t^{ngr} + \nu_t).
\end{aligned} \tag{4}$$

[20] Despite being quite restrictive, the assumption that the effect of exceptional events is constant over time does not affect our results but conveniently reduces notation.

In model (3) the coefficient of interest is β, whose components measure the average response of sovereign risk in each country to a shock in Greece. We cannot however use a standard Ordinary Least Squares (OLS) technique to estimate β as we cannot sustain the assumption of conditional homoskedasticity in $\tilde{u}_t$. By construction, u_t^{gr} from (4) depends on the "normal" shocks ν_t only, while u_t^{ngr} is the sum of ν_t and, potentially, the "exceptional" shocks $d_{i,t}\xi_t^i$ in all countries. Our strategy thus forces us to derive an estimator that allows for the presence of two alternating regimes in the variance-covariance matrix of the residuals.

Case 1: Two Regimes of VCV. Writing $\tilde{\Phi} = [\Phi, \beta]$ and $\tilde{z}_t = [z_t', \mathbf{1}_{gr,t}]'$, we can reformulate model (3) as

$$y_t = \tilde{\Phi}\tilde{z}_t + \tilde{u}_t,$$

which is formally equivalent to (1). Accounting for heteroskedasticity, we assume that

$$E_t(\tilde{u}_t) = 0, \qquad E_t(\tilde{u}_t\tilde{u}_t') = \begin{cases} \Sigma_{gr} & \text{if } t \in \mathcal{T}^{gr} \\ \Sigma_{ngr} & \text{otherwise} \end{cases}, \qquad E_t(\tilde{u}_t\tilde{u}_s') = 0 \text{ for } s \neq t,$$

where $\mathcal{T}^{gr}$ is the known subset of pure Greek events and the two matrices Σ_{gr} and Σ_{ngr} have full rank and are invertible. We refer to the Appendix for the derivation of a generalized least squares (GLS) estimator of $\tilde{\Phi}$, which accounts for this specific type of heteroskedasticity.

In addition to the arising heteroskedasticity, we need to address another complication of the narrative approach. Specifically, we cannot infer the rate of cross-country transmission of sovereign risk directly from our estimates of β. By construction each component β_k equals the actual rate of transmission from Greece to country k, as measured by the component $B_{(k,gr)}$, times the average magnitude of Greek events ξ^{gr}. To obviate this problem, we exploit the linearity of our model and estimate the rate of transmission by looking at the response of each country k relative to that of Greece, or β_k/β_{gr}. The coefficient β_{gr} equals ξ_t^{gr}, so this ratio captures exactly $B_{(k,gr)}$. We then compute a confidence interval for each of these ratios by bootstrapping the residuals from regression.[21]

[21] In doing so, we take special care to resample the residuals only within each variance-covariance regime, not across them, to satisfy the assumption of invariance in distribution that underlies the bootstrapping technique.

3.2 Inspecting the Role of Financial Linkages

In order to identify the contribution of financial linkages to the overall transmission of sovereign risk, we exploit the time variation in the exposure of other countries to Greek debt. This difference-in-difference approach enables us to disentangle the role of financial linkages from other channels of transmission, ranging from trade interdependence to non-fundamental-based linkages. We observe, indeed, that financial exposure tends to move much faster than real bilateral linkages, and its evolution is not directly linked to changes in market sentiments.

Specifically, we allow the spillover matrix B in (2) to vary over time, reflecting a change in the financial linkages across countries. As our measure of financial linkages is reported on a quarterly basis, we let B_q to denote the transmission matrix in quarter q and we assume

$$B_q = B_0 + dL_q, \tag{5}$$

where B_0 is a constant matrix capturing all time-invariant channels and the components of L_q are our measures of financial linkages, which vary across quarters. In (5), d measures the effect of a change in financial linkages on the transmission of sovereign risk and is assumed to be equal across country pairs.[22]

Substituting for (5) into (2), we can rewrite our baseline model as

$$y_t = \Phi z_t + \left(\delta_0 + \delta_1 l_{gr,q}\right) \cdot \mathbf{1}_{gr,t} + \tilde{u}_t, \tag{6}$$

where $\delta_0 \equiv (I + B_0)\xi^{gr}$ is a $n \times 1$ vector, $\delta_1 \equiv d \cdot \xi_{gr}$ is a scalar and $l_{gr,q}$ is a $n \times 1$ vector of cross-border financial exposures to Greece.

Our main goal is to estimate the coefficient δ_1 in (6), from which we will subsequently obtain the value of the parameter of interest d through divsion by the coefficient measuring the Greek average response. Technically, δ_1 is assumed to be constant across the equation of the VAR model and therefore its estimation requires a cross-equation restriction on the interaction coefficient appearing in the above specification. Furthermore, implicit in model (6) is the fact that the residual $\tilde{u}_t$ now depends on B_q (see (4)) and thus has a variance-covariance matrix that can vary both across regimes (Greek and non-Greek) and across

[22]Furthermore, we maintain our previous normalization and set all diagonal entries of B_0 to one (the diagonal entries of L_q are zero by construction).

quarters q. We thus need to derive an estimator that allows us to address these properties of the model.

Case 2: Restricted Regressors and N Regimes of VCV. Consider the following model

$$y_t = \tilde{\Phi}\tilde{z}_t + \tilde{u}_t \qquad\qquad \tilde{\Phi}'_{(k)} = R_k \cdot c_k, \tag{7}$$

for $k = 1, \ldots, n$. Here, $\tilde{\Phi} = [\Phi, \delta_0, \Delta_1]$, $\tilde{z}_t = [z_t', \mathbf{1}_{gr,t}, l'_{gr,q} \cdot \mathbf{1}_{gr,t}]'$. Thus, Δ_1 denotes a vector of coefficients on the interaction term $l_{gr,q} \cdot \mathbf{1}_{gr,t}$. The equation on the right specifies a set of linear restrictions on each row $\tilde{\Phi}_{(k)}$ of the matrix of coefficients $\tilde{\Phi}$. Imposing

$$R_k = \begin{pmatrix} I_{p\times p} & 0_{p\times 1} & 0_{p\times 1} \\ 0_{1\times p} & 1 & 0 \\ 0_{n\times p} & 0_{n\times 1} & e^k \end{pmatrix} \qquad \text{and} \quad c_k = [\Phi_{(k)}, \delta_{0,k}, \delta_1]'$$

where e^k is the k^{th} unit vector of length n, this model is identical to our specification (6).

The residual process satisfies the following set of assumptions:

$$E_t(\tilde{u}_t) = 0, \qquad E_t(\tilde{u}_t\tilde{u}_t') = \begin{cases} \Sigma_{gr,q} & \text{if } t \in \mathcal{T}_q \cap \mathcal{T}^{gr} \\ \Sigma_{ngr,q} & \text{if } t \in \mathcal{T}_q \backslash \mathcal{T}^{gr} \end{cases}, \qquad E_t(\tilde{u}_t\tilde{u}_s') = 0 \text{ for } s \neq t, \tag{8}$$

where $\mathcal{T}_q$ is the set of dates in quarter q and $\mathcal{T}^{gr}$ is defined as in Case 1. This formulation allows the variance-covariance matrix of residuals to depend on the quarter q within the two regimes. We refer once again to the Appendix for a technical description of the restricted GLS estimator we use to estimate model (7) under the specific type of heteroskedasticity assumed in (8).

We are now ready to perform our estimations. Specifically, in a two-step procedure, we would first estimate (6) by simple OLS, compute the residuals to obtain estimates for the quarter-specific VCV matrices $\Sigma_{ngr,q}$ and $\Sigma_{gr,q}$ and then use these matrices to estimate the feasible GLS. The value of d can then be inferred by normalizing the estimate of δ_1 by the average size of Greek events, which, by our normalization of B_0, corresponds to the Greek component of δ_0. We then construct the confidence interval by repeating the two-step estimation procedure with bootstrapped residuals to obtain a synthetic distribution for the estimate of d.

Continuing on this path, however, we face the difficulty that the quarterly sets of dates with Greek events, $\mathcal{T}_q$, contain, on average, less than two elements. Thus, the standard estimate of the quarter-specific variance-covariance matrix

$$\hat{\Sigma}_{gr,q} = \frac{1}{|\mathcal{T}_q \cap \mathcal{T}^{gr}|} \sum_{t \in \mathcal{T}_q \cap \mathcal{T}^{gr}} \tilde{u}_t \tilde{u}_t'$$

delivers matrices that do not have full rank and cannot be inverted for the calculation of the GLS estimations. We address this problem by simply taking the time-average variance-covariance matrix for all quarters.

4 Results

This section summarizes the results of the GLS estimations of the transmission of sovereign risk. We begin by estimating model (3) to assess average, time-invariant transmission. We then turn to model (6) to analyze the role of the financial linkages.

Unless explicitly described otherwise, the dependent vector y_t consists of five-year sovereign CDS premia of eleven European countries (including Greece), logged and differentiated.[23] We take logs to avoid giving excessive weight to the relatively large reactions of "PIGS' or other troubled countries (compare Figure 1). Intuitively, a jump in the German CDS premium from 30 to 50 basis points should be considered much more dramatic than an increase of the Portuguese CDS premium from 450 to 470 basis points. The vector of independent variables $\tilde{z}_t$ comprises the lags y_t, exogenous variables x_t, the Greek dummy and, when estimating model (6), the interaction of the Greek dummy with a measure of financial exposure to Greece. To control for common shocks affecting global market conditions, the vector of exogenous variables consists of the following variables: sovereign CDS premia of the US and Japan (logged and differentiated as all other CDS premia), the VIX index for the US (logged), stock market returns for the US (defined as log changes of the S&P index), the US federal fund rate, and seasonal dummies for each day of the week.

The data are reported on a daily basis, except for the data on financial linkages. They span a period of 500 business days for the calendar years 2009 and 2010.[24]

[23]We differentiate the CDSs to make them stationary. The estimation results remain intact for HP-filtered data.

[24]The eleven countries are Austria, Belgium, France, Germany, Greece, Italy, the Netherlands, Portugal, Spain, Sweden and UK.

4.1 Baseline Specification Time-Invariant Transmission

In our baseline specification, we estimate model (3). The coefficient β is constant over the entire period and thus captures the time-average of the responses to the identified Greek shocks by the ten other European countries.

Our tests of autocorrelation of the residuals prompt us to include eight lags of the dependent variable and no lag of the exogenous variables. Moreover, we disregard heteroskedasticity within the set of days where Greek events occur.

Table 4a reports the results for the baseline specification. The columns correspond to the elements of the vector y_t, i.e., to the eleven countries. In the upper panel, we report the corresponding coefficients of the indicator for Greek events (GR), along with the standard errors. All of the estimates of the eleven coefficients are positive, and the implied t-ratios indicate that the coefficients are statistically significant on the five-percent level for more than half of the countries. Increasing the confidence level to ten percent, only Germany and the Netherlands show non-significant responses to Greek shocks. The point estimate for Greece itself is 0.0631, which indicates that the events identified in Table A1 generated, on average, a jump in the Greek sovereign CDS premium of over 6 percent. At a time-average of 679 basis points in 2010, this average change in the CDS premium is roughly equivalent to a jump of 43 basis points ($679 * .0631 = 42.8449$).

The coefficients corresponding to countries other than Greece itself vary from 0.0036 (Germany) to 0.0277 (Italy). Measured relative to the magnitude of the Greek coefficient (lower panel), these estimates imply a rate of transmission that ranges from 5.71% to 43.90%, with an average for all ten countries equal to 28.43%.[25] These numbers indicate economically important transmission rates. At the same time, the estimated coefficients are strikingly similar in magnitude. To put this observation into perspective, recall that the CDS premia are logged so that the coefficients measure the transmission of Greek shocks in terms of percentage change in each countrys sovereign CDS premium. Thus, a one-percent increase in the Portuguese CDS premium corresponds to a jump of 2.9 basis points, whereas a one-percent increase in the French CDS premium corresponds to an increase of 0.7 basis points based on 2010 averages.

[25] Notice that there is a strong correspondence between the statistical significance of the normalized and of the absolute responses. The only exception is the UK, where we observe a slight loss of significance. We attribute this difference to the use of bootstrap confidence intervals in the case of normalized response.

Assessing transmission rates by levels, Table 4b repeats the estimates without prior logging the CDS premia. Again, all estimates are positive (except for Germany, where the estimate is slightly negative), although the significance drops for Spain and Portugal. The coefficient for Greece indicates an average jump of 30.15 basis points in response to Greek events, which is somewhat lower than, but in the same realm as, the benchmark computed using the elasticities. The responses of European sovereign CDS premia relative to the response of the Greek CDS premium ranges between −0.04% and 16.63%.

Figure 3 plots the impulse response functions for all eleven countries, including the five-percent confidence bounds. The figure shows that a Greek shock generates a positive contemporaneous response in the CDS premium of all countries, whose significance is comparable to the levels reported in Table 4a. In general, the response becomes insignificant after one day. More importantly, the figure shows no negative response as the lag increases, which indicates that the effects of Greek shocks are not transitory and the sovereign CDS premia do not revert to their initial levels over time.[26]

Overall, the results from the baseline regression indicate substantial rates of transmission of sovereign risk across Europe. In the next step, we analyze the role of financial linkages in that transmission.

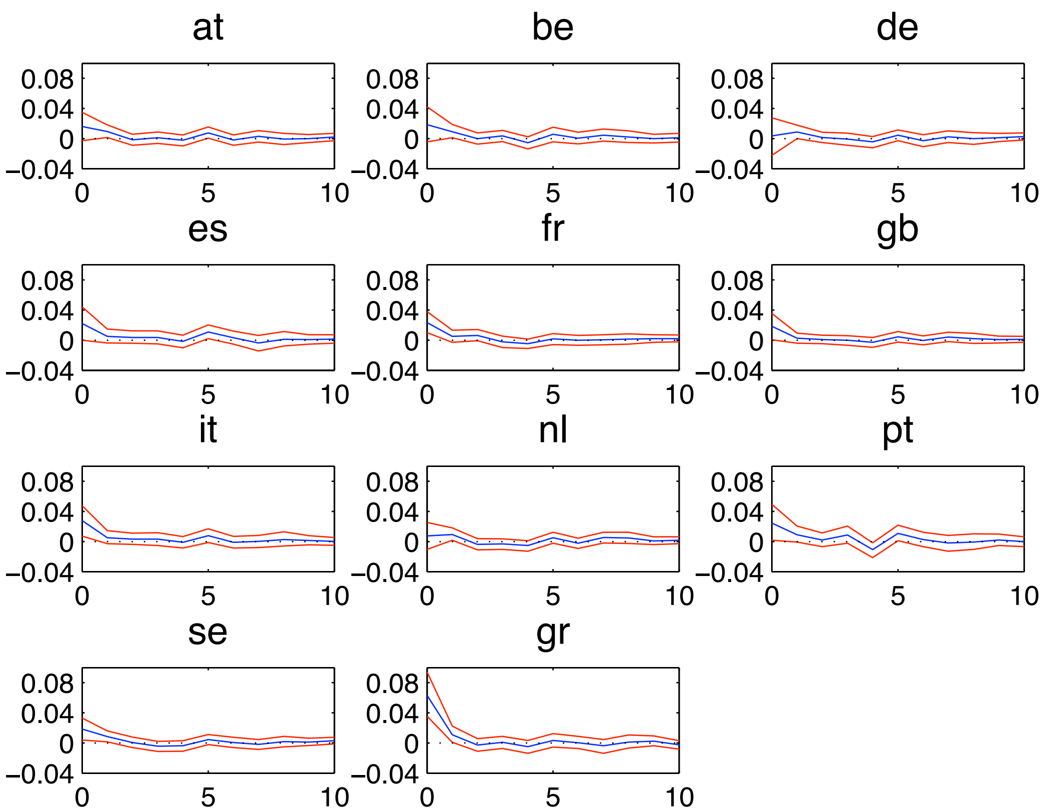
4.2 Financial Linkages

We turn now to the estimation specified in (6) to analyze the role of financial linkages in the Euro crisis, exploiting the time-variation of the exposure to Greek debt. We use three measures for the linkages: exposure (i) to Total Greek Debt; (ii) to Greek Public Debt; and (iii) to Debt of Greek Banks. All measures are normalized by average real GDP of the years 2009 and 2010 and are logged. Whenever there is no risk of confusion, we will refer to these measures as *exposure to Greece*.

Table 5 reports the estimation results based on model (6). Column I corresponds to the specification that relies on exposure to Greece defined as Total Greek Debt. The first row of the table reports the coefficients on the Greek dummy (GR), measuring the Greek fixed effect of the average Greek shock. The coefficient is 0.0679, which is equal

[26] The confidence bounds are derived from a bootstrap distribution based on 1000 replications.

Figure 3: Impulse Response Functions



Impulse response functions of CDS (log-changes) of eleven countries to a Greek shock of average size for the 10 business days after impact (blue line). The red lines indicate the 5% confidence bounds based on a bootstrap exercise with 1000 replications.

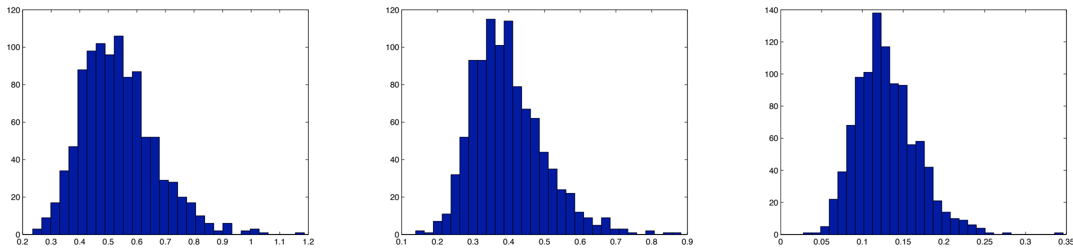
to the point estimate in our baseline regression (Table 4a). As for the fixed effects on other countries, we choose to not report the estimates, although we check that the implied average rate of transmission is consistent with the baseline results. In our analysis of the role of financial linkages in the transmission of sovereign risk, the coefficient of interest is the one on $GR * Total$, i.e., the interaction term between the Greek dummy and the exposure to Greece ($l_{gr,q} \cdot \mathbf{1}_{gr,t}$ in (6)). This estimated coefficient is positive and significant at the one-percent level, as indicated by the implied t-ratio.

To assess transmission independent of the size of the original shock, the coefficient of $GR * Total$ needs to be normalized by the average size of the Greek shock. Because we cannot observe the magnitude of the underlying shock, we normalize the coefficient on $GR * Total$ by the Greek component of the Greek dummy (GR). The resulting ratio measures how an increase in exposure to Greek debt affects the rate of transmission and thus corresponds to the coefficient d in (5). The point estimates in Column I result in a value of 0.5538 for d. As discussed in Section 3, we cannot derive a theoretical distribution for the estimator of d and resume bootstrapping to assess statistical significance. The lower part of Table 4 summarizes the bootstrap confidence bounds based on 1000 bootstrap replications, indicating that the one-percent confidence interval is on the positive axis. Based on a one-percent confidence level, we can thus conclude that the bilateral exposure to Greek debt plays a positive role in the transmission of sovereign risk. The exposure to Greek debt is not only statistically significant but also economically relevant: the value of 0.5538 for d indicates that a one-percent reduction in the exposure to Greek debt decreases a country's CDS response to a Greek shock by more than one half of one percent. Based on a rough calculation, the transmission rate to the country with the greatest average exposure to Greece (1.22 percent of GDP) is, everything else equal, 46 percent larger than the transmission rate to the country with the lowest average exposure (0.08 percent of GDP). We take these results as motivation to scrutinize the different sub-components of Total Greek Debt, i.e., Greek Public Debt and Debt of Greek Banks.

Column II corresponds to the specification where exposure to Greece is defined by Greek Public Debt. The coefficient of the dummy GR is 0.0675, and the coefficient of the interaction term $GR * Public$ equals 0.0266. The implied t-ratios of both coefficients indicate significance at the one-percent level. The ratio of the two point estimates in Column I is 0.3941 for d. The lower part of Column II shows that the bootstrap confidence interval for the one-percent confidence level is on the positive axis. In addition, the value of d lies at the center of the confidence intervals and is close to the mean of the sample of bootstrap replications ($mean$ ($bootstr$)), which is 0.4022. Figure 5(b) illustrates the bootstrap distribution for d, the coefficient indicating transmission through financial linkages.

The value of 0.3941 for d indicates that a one-percent reduction in exposure to Greek public debt decreases a country's CDS response to a Greek shock by more than one third of one percent. Combining this information with an average reduction of 27% in the exposure to Greek public debt (compare Table 2b), our results indicate that the response to Greek

Figure 4: Bootstrap Distribution of the Rate of Response to Greek News



(a) Total Debt (b) Public Debt (c) Bank Debt

Distribution of the estimated parameter of the rate of response (d in Tables 3-8), derived from 1000 bootstrap replications. The panels correspond to the financial exposure to Greek Total Debt, Greek Public Debt and Debt of Greek Banks.

shocks might have declined as a consequence of the decrease of foreign exposure by roughly ten percentage points ($0.3941 * 0.27 = 0.1064$), which correspond to more than one third of the average rate of transmission estimated in Section 4.1, i.e., 28.43%.

Column III of Table 5 reports parallel results when exposure to Greece is defined by Debt of Greek Banks. Again, the coefficient of the interaction term $GR * Bank$ is positive and significant at the one-percent level, and the implied value of d is 0.1469. The bootstrap confidence bounds indicate that the estimate of d is significant at the one-percent level. The magnitude of d suggests that the role of exposure to debt of foreign banks is somewhat less important for the transmission of sovereign risk: the elasticity of the spillover effects to bank debt exposure is about one third of the corresponding elasticity with respect to direct exposure to Greek sovereign debt.

The value of 0.1469 for the rate of transmission through interbank lending indicates that a one-percent reduction in the exposure to Greek banks decreases a country's CDS response to a Greek shock by almost 0.15 percent. Again, in a rough calculation, the relative response of the sovereign default risk of other European countries to a Greek shock would have been 15 percent smaller in the absence of any financial exposure.

Overall, the baseline estimates suggest that the rates of exposure to Greek sovereign debt and to debt of Greek banks played important roles in the spreading of sovereign risk across the Euro area. Interestingly, our estimate of the elasticity of CDS responses to changes in Greek Total Debt (0.5538) appears to decompose linearly into a bit less than one third stemming from exposure to Debt of Greek Banks (0.1469) and more than two thirds stemming from exposure to Greek Public Debt (0.3941).[27]

4.3 Robustness

We conduct a number of robustness checks. First, we repeat our estimations measuring CDS premia in basis points and the exposure to Greek debt as a percentage of GDP (i.e., we do not log these variables). In this linear specification, the CDS premia of the European countries are assumed to react to a Greek shock independent of their respective levels. An increase of the German CDS premium from 30 to 50 basis points is thus treated similarly to an increase of the Portuguese CDS premium from 450 to 470 basis points. Moreover, an increase of GDP-normalized exposure from 0.01 to 0.02 is assumed to have the same effect as an increase from 0.11 to 0.12. This specification therefore puts less weight on the time variation of very lightly exposed countries. Table 6 shows that both the sign and the significance of the estimates are preserved under this linear specification. We observe, however, that the estimate of the coefficient d in the specification involving the Debt of Greek Banks is now larger than the estimate in the specification using Greek Public Debt. The estimates of the coefficient d are 12.19 and 40.05, respectively. The reverse order in the magnitude of d reflects the fact that the exposure to bank debt, which equals, on average, 0.09% of GDP, is significantly lower than the average exposure to sovereign debt (0.48%). With these point estimates, based on exposure as a percentage of GDP, we can give the following absolute interpretation of our results. The value for the relative response of 12.19 suggests that a reduction of exposure to Greek Public Debt of 0.1 percent of GDP (the average exposure is 0.48 percent of GDP) implies that the response to Greek shocks decreases by 0.012 basis points ($0.001 * 12.19 = 0.012$). Like the estimates with the logged variables, these results indicate that financial linkages are important for the transmission of sovereign risk.

Next, we limit the number of countries in our sample. Specifically, we exclude Sweden and the UK from our sample, which limits the analysis to countries within the Euro zone. We thus include only the countries that assume implicit liabilities through the ECB balance

[27] It is clear that this calculation is overly simplified because the elasticity of variable y to variable x does not linearly decompose into the elasticities of y to the subcomponents of x.

sheet. A comparison between the results of the reduced and the full samples can therefore indicate whether the transmission of sovereign risk operates significantly through ECB assets, in which case the coefficient d could be biased. Table 7 shows that the elasticity of transmission to a change in financial linkages, d, is positive and significant on the one-percent level for all specifications. Moreover, for the reduced sample, all three estimates of d are very close to those of the full sample (compare Table 5). We can thus conclude that our results are not driven by the specific sample of countries considered.

We then estimate our baseline specification, replacing the contemporaneous exposure to Greek debt with the lagged exposure. This specification accounts for possible delays in the diffusion of information regarding cross-border exposures, in which case the relevant exposure for the transmission of sovereign risk is the one observed in the quarter ahead of the date of a Greek shock. The results are shown in Table 8. The estimates for the coefficient of interest, d, are all positive and significant at the one-percent level. Furthermore, the point estimates are very close to those estimated in the specification with contemporaneous exposure, with the only exception being the case of Public Debt, in which the value of d drops by one third. Overall, we can conclude that the results appear robust to considerations about the speed of information diffusion.

As an additional robustness check, we try to capture not only the spillover effects of direct exposure to Greek debt, but the entire network of financial linkages across the countries in our sample. In particular, if French banks are strongly exposed to Greek sovereign debt but Spanish banks are not, the Spanish banks might nevertheless suffer from an increase in Greek sovereign risk due to their indirect exposure through the French banking system. To capture these indirect effects, we invert the matrix of bilateral financial linkages and estimate the corresponding Greek column of the matrix of financial linkages. The exposure now captures the overall – i.e., the direct and the indirect – transmission of Greek shocks to the respective European countries. Table 9 shows that the estimates for d are significant at the one-percent level in all three specifications. Interestingly, the point estimates are larger than those reported in Table 5, with an increase that ranges between one fourth in the specification with Greek Public Debt and three fourths in the specification with Debt of Greek Banks. These numbers demonstrate the importance of accounting for the entire network of cross-border financial linkages when studying the propagation of sovereign risk, particularly in the context of the highly integrated European countries.

Our final concern is that days of Greek news might be interpreted by the market as days of deteriorating global conditions, which would then be reflected in larger CDS premia independent of the direct linkages of countries with Greece. Under this hypothesis, the response to Greek shocks would thus correlate more strongly with the change in the general exposure to risk than with the change in exposure to Greek debt. We thus re-estimate our model, replacing our three measures of exposure to Greek debt with the corresponding exposures to foreign debt, defined as the sum of the claims on each foreign country and for each class of assets. The results are reported in Table 10. The point estimates for d in the specification with Total Foreign Debt and Debt of Foreign Banks still have a positive sign, but their significance is strongly rejected by the sample of bootstrap replications, which is now centered around zero. For the specification with Foreign Public Debt, the estimates for d are now negative and significant at the five-percent level. Overall, these results indicate that exposure to foreign debt from origins besides Greece is at best unrelated to the reaction of sovereign risk, and in one case the relation is the inverse of the one expected.

5 Conclusion

This paper has shown that financial linkages matter for the transmission of sovereign risk in the Euro crisis. Our estimates show that a 10-percent increase in the exposure to foreign sovereign debt increases the spillover effects of sovereign risk by 1.8 percent. Similarly, a 10-percent increase in exposure to debt of foreign banks increases these spillover effects by 1.2 percent. These estimates are statistically significant and economically relevant.

Methodologically, we follow the narrative approach of Romer and Romer (1989), identifying financially relevant news shocks that can be attributed to Greece's problems servicing its sovereign debt. These shocks are used to assess the response of sovereign risk of other European countries. We further relate the latter responses to cross-border financial linkages, controlling for transmission channels that are slow to change through fixed effects.

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A The Estimator

We derive the estimator used in main body of the paper. To do so, we will proceed in several steps. First, we state a technical result on matrix inversion, which we use later. Second, we derive a multi-variate generalized least squares (GLS) estimator under the type of heteroskedasticity described in the main text. Third, we describe our bootstrapping techniques.

A.1 Matrix Algebra

Claim 1 *Let Σ_k be a collection of $n \times n$ matrices with full rank, for $k = 1, ..., K$. Let further I_k denote $m \times m$ matrices of the following form: all off-diagonal elements equal to zero, all diagonal elements are either zero or one and the I_k sum to the unit matrix: $\sum_k I_k = \mathbf{1}_m$. Then,*

$$\left(\sum_{k=1}^{K} \Sigma_k \otimes I_k \right)^{-1} = \sum_{k=1}^{K} \Sigma_k^{-1} \otimes I_k \tag{9}$$

holds, where $\otimes$ symbolizes the Kronecker matrix multiplication.

Proof. Multiplication of $\sum_k \Sigma_k \otimes I_k$ and the right hand side in (9) and exploiting the basic properties of the Kronecker multiplication yields

$$\begin{aligned}
\left(\sum_k \Sigma_k \otimes I_k \right) \sum_k \Sigma_k^{-1} \otimes I_k &= \sum_k \left(\Sigma_k \otimes I_k \right) \left(\Sigma_k^{-1} \otimes I_k \right) + \sum_{k \neq j} \left(\Sigma_k \otimes I_k \right) \left(\Sigma_j^{-1} \otimes I_j \right) \\
&= \sum_k \Sigma_k \Sigma_k^{-1} \otimes I_k I_k + \sum_{k \neq j} \Sigma_k \Sigma_j^{-1} \otimes I_k I_j \\
&= \sum_k \mathbf{1}_n \otimes I_k + \sum_{k \neq j} \Sigma_k \Sigma_j^{-1} \otimes \mathbf{0}_m \\
&= \mathbf{1}_n \otimes \sum_k I_k = \mathbf{1}_{nm}
\end{aligned}$$

■

A.2 GLS Estimator – Two Regimes of VCV-Matrix

Consider the n-dimensional model

$$y_t = \tilde{\Phi} \tilde{z}_t + \tilde{u}_t, \tag{10}$$

where $\tilde{\Phi} = [\Phi; \beta]$ and $\tilde{z}_t = [z_t'; \mathbf{1}_{gr,t}]'$. The residuals $\tilde{u}_t$ satisfy

$$E_t(\tilde{u}_t) = 0, \qquad E_t(\tilde{u}_t\tilde{u}_t') = \begin{cases} \Sigma_{gr} & \text{if } t \in \mathcal{T}^{gr} \\ \Sigma_{ngr} & \text{if } t \notin \mathcal{T}^{gr} \end{cases}, \qquad E_t(\tilde{u}_t\tilde{u}_s') = 0 \text{ for } s \neq t \tag{11}$$

where $\mathcal{T}^{gr}$ is the set of greek events and Σ_{gr} and Σ_{ngr} are assumed to be non-singular. Here and in the following, E_t denotes the expected value given the information set available at period t.

Further assumptions are required, in particular, to guarantee the asymptotic normality of the GLS estimator. Following Lütkepohl (2007, p. 397), we assume throughout that the residual $\tilde{u}_t$ is a white noise process, the matrix $\tilde{\Phi}$ of autoregressive coefficients satisfies a stability condition, and that the vector of exogenous variables x_t, which appears among the elements of z_t, is generated by a stationary, stable VAR process which is independent of the process $\tilde{u}_t$.

For a sample size T, model (10) can be written compactly as

$$\mathbf{y} = (I_n \otimes \tilde{Z}')\tilde{\boldsymbol{\phi}} + \tilde{\mathbf{u}}, \tag{12}$$

where $\mathbf{y}$ is the $nT \times 1$ vector obtained by stacking the T-dimensional vectors $y_{(k)}$ formed by the time series of each k^{th} component of y_t, $\tilde{\mathbf{u}}$ is similarly defined by stacking the vectors $u_{(k)}$ formed by the time series of the components of $\tilde{u}_t$, I_n is a n-dimensional identity matrix and $\tilde{Z} = [\tilde{z}_1, \ldots, \tilde{z}_T]$ is the $(p+1) \times T$ matrix formed by the time series of vector $\tilde{z}_t$. Here, $p+1$ denotes the number of regressors in each equation of model (10), which include $p = nJ + m$ variables plus the Greek dummy. $\tilde{\boldsymbol{\phi}} = [\tilde{\Phi}_{(1)}', \tilde{\Phi}_{(2)}', \ldots, \tilde{\Phi}_{(n)}']'$ is the $n(p+1) \times 1$ vector obtained from stacking the rows $\tilde{\Phi}_{(k)}$ of matrix $\tilde{\Phi}$.

Consistently with the assumptions in (11), the variance-covariance of the vector of residuals $\tilde{u}_t$ satisfies

$$\Sigma_{\tilde{\mathbf{u}}} = \Sigma_{gr} \otimes I_T^{gr} + \Sigma_{ngr} \otimes I_T^{ngr}.$$

Here I_T^{gr} is a T-dimensional diagonal matrix whose diagonal entries corresponds to the entries of the vector $[\mathbf{1}_{gr,1}, \ldots, \mathbf{1}_{gr,T}]$ and $I_T^{ngr} = I_T - I_T^{gr}$. The GLS estimator of $\tilde{\boldsymbol{\phi}}$ thus minimizes $\tilde{\mathbf{u}}'(\Sigma_{\tilde{\mathbf{u}}})^{-1}\tilde{\mathbf{u}}$ and equals

$$\widehat{\tilde{\boldsymbol{\phi}}} = [(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')]^{-1}(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}\mathbf{y}. \tag{13}$$

Under the assumption in (11) and the regularity conditions discussed above, we can show

Claim 2 *The GLS estimator of $\tilde{\beta}$ converges in distribution according to*

$$\sqrt{T}\left(\widehat{\tilde{\boldsymbol{\phi}}} - \tilde{\boldsymbol{\phi}}\right) \xrightarrow{d} N(\mathbf{0}, \mathcal{Q}^{-1}\mathcal{V}\mathcal{Q}^{-1}),$$

where $\mathcal{Q} \equiv plim\left(T^{-1}(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')\right)$ is well-defined, symmetric and non-singular and $\mathcal{V} \equiv (I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')$.

Proof. Rearranging (13) we obtain

$$\sqrt{T}\left(\widehat{\tilde{\boldsymbol{\phi}}} - \tilde{\boldsymbol{\phi}}\right) = \left[\frac{(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')}{T}\right]^{-1} \frac{(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}\tilde{\mathbf{u}}}{\sqrt{T}}. \tag{14}$$

Expanding the term within squared brackets yields

$$\frac{(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')}{T} = \frac{1}{T}\begin{pmatrix} \tilde{Z}\Omega_{11}\tilde{Z}' & \tilde{Z}\Omega_{12}\tilde{Z}' & \dots & \tilde{Z}\Omega_{1n}\tilde{Z}' \\ \tilde{Z}\Omega_{21}\tilde{Z}' & \tilde{Z}\Omega_{22}\tilde{Z}' & \dots & \tilde{Z}\Omega_{2n}\tilde{Z}' \\ \vdots & \vdots & \ddots & \vdots \\ \tilde{Z}\Omega_{n1}\tilde{Z}' & \tilde{Z}\Omega_{n2}\tilde{Z}' & \dots & \tilde{Z}\Omega_{nn}\tilde{Z}' \end{pmatrix}$$

where Ω_{ij} denotes the $T \times T$ block of $\Sigma_{\tilde{\mathbf{u}}}^{-1}$ with (i, j) position. Applying Claim 1, we write

$$\Omega_{ij} = \omega_{ij,gr} \cdot I_T^{gr} + \omega_{ij,ngr} \cdot I_T^{ngr},$$

where $\omega_{ij,gr}$ denotes the component of Σ_{gr}^{-1} with (i, j) position. $\omega_{ij,ngr}$ is defined analogously. Substituting for Ω_{ij} we obtain

$$\frac{1}{T}\tilde{Z}\Omega_{ij}\tilde{Z}' = \omega_{ij,gr}\lambda \sum_{t\in\mathcal{T}^{gr}} \frac{\tilde{z}_t\tilde{z}_t'}{T_{gr}} + \omega_{ij,ngr}(1-\lambda)\sum_{t\notin\mathcal{T}^{gr}} \frac{\tilde{z}_t\tilde{z}_t'}{T_{ngr}},$$

where T_{gr} and T_{ngr} are used to denote the number of Greek and non-Greek events respectively and $\lambda = \frac{T_{gr}}{T}$ is the sample probability of $t \in \mathcal{T}^{gr}$. Assuming that λ converges to a finite number as $T \to \infty$, the assumptions in (11) and the regularity conditions imposed above ensure that each $\tilde{Z}\Omega_{ij}\tilde{Z}'$ has a well-defined probability limit. We can thus write

$$\frac{(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')}{T} \xrightarrow{p} \mathcal{Q}, \tag{15}$$

where $\mathcal{Q}$ is a symmetric matrix (since $\Sigma_{\tilde{\mathbf{u}}}^{-1}$ is symmetric) and is non-singular.

Expanding the second term in (14) we get

$$\frac{(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}\tilde{\mathbf{u}}}{\sqrt{T}} = \frac{1}{\sqrt{T}}\begin{pmatrix} \sum_{j=1}^{n} \tilde{Z}\Omega_{1,j}\tilde{u}_{(j)} \\ \sum_{j=1}^{n} \tilde{Z}\Omega_{2,j}\tilde{u}_{(j)} \\ \vdots \\ \sum_{j=1}^{n} \tilde{Z}\Omega_{n,j}\tilde{u}_{(j)} \end{pmatrix},$$

where

$$\frac{1}{\sqrt{T}}\sum_{j=1}^{n}\tilde{Z}\Omega_{i,j}\tilde{u}_{(j)} = \sum_{j=1}^{n}\left(\omega_{ij,gr}\sqrt{\lambda}\sum_{t\in\mathcal{T}^{gr}}\frac{\tilde{z}_t\tilde{u}_{j,t}}{\sqrt{T_{gr}}} + \omega_{1j,ngr}\sqrt{1-\lambda}\sum_{t\notin\mathcal{T}^{gr}}\frac{\tilde{z}_t\tilde{u}_{j,t}}{\sqrt{T_{ngr}}}\right).$$

The terms $\tilde{z}_t\tilde{u}_{j,t}$ are martingale difference sequences with well-defined variance-covariance matrices in both Greek and non-Greek regimes. We can therefore apply a version of the Central Limit Theorem (see Hamilton (1994, p. 193)) to show that each $\frac{1}{\sqrt{T}}\sum_{j=1}^{n}\tilde{Z}\Omega_{i,j}\tilde{u}_{(j)}$ has a well-defined asymptotic normal distribution. Thus, we can write

$$\frac{(I_n\otimes\tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}\tilde{\mathbf{u}}}{\sqrt{T}}\xrightarrow{d}N(\mathbf{0},\mathcal{V}), \tag{16}$$

where $\mathcal{V}\equiv(I_n\otimes\tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n\otimes\tilde{Z}')$. Combining (15) and (16), we obtain

$$\sqrt{T}\left(\widehat{\tilde{\boldsymbol{\phi}}}-\tilde{\boldsymbol{\phi}}\right)\xrightarrow{d}N(\mathbf{0},\mathcal{Q}^{-1}\mathcal{V}\mathcal{Q}^{-1}),$$

■

A.3 GLS Estimator – Time-Dependent VCV-Matrix and Constraints

Consider the following model

$$y_t=\tilde{\Phi}\tilde{z}_t+\tilde{u}_t \qquad\qquad \tilde{\Phi}'_{(k)}=R_k\cdot c_k, \tag{17}$$

for $k=1,\dots,n$. Here $\tilde{\Phi}=[\Phi,\delta_0,\Delta_1]$, $\tilde{z}_t=[z_t',\mathbf{1}_{gr,t},l_{gr,q}'\cdot\mathbf{1}_{gr,t}]'$. Thus, Δ_1 denotes a vector of coefficients on the interaction term $l_{gr,q}\cdot\mathbf{1}_{gr,t}$. The equation on the right specifies a set of linear restrictions on each row $\tilde{\Phi}_{(k)}$ of the matrix of coefficients $\tilde{\Phi}$. In particular,

$$R_k=\begin{pmatrix} I_{p\times p} & 0_{p\times 1} & 0_{p\times 1}\\ 0_{1\times p} & 1 & 0\\ 0_{n\times p} & 0_{n\times 1} & e^k\end{pmatrix} \qquad \text{and} \quad c_k=[\Phi_{(k)},\delta_{0,k},\delta_1]'$$

where e^k is the k^{th} unit vector of length n. This set of constraints restricts the components of vector Δ_1 to be equal to each other.

The residual process satisfies the following set of assumptions

$$E_t(\tilde{u}_t)=0, \qquad E_t(\tilde{u}_t\tilde{u}_t')=\begin{cases}\Sigma_{gr,q} & \text{if } t\in\mathcal{T}_q\cap\mathcal{T}^{gr}\\ \Sigma_{ngr,q} & \text{if } t\in\mathcal{T}_q\backslash\mathcal{T}^{gr}\end{cases}, \qquad E_t(\tilde{u}_t\tilde{u}_s')=0 \text{ for } s\neq t, \tag{18}$$

where $\mathcal{T}_q$ is the set of dates in quarter q and $\mathcal{T}^{gr}$ is defined as before. This formulation allows the variance-covariance matrix of residuals to depend on the quarter q within the two regimes.

Rewriting model (17) using the notation introduced in the previous case, we obtain

$$\mathbf{y} = (I_n \otimes \tilde{\mathbb{Z}}')\tilde{\boldsymbol{\phi}} + \tilde{\mathbf{u}} \qquad\qquad \tilde{\boldsymbol{\phi}} = R \cdot \boldsymbol{\gamma}. \tag{19}$$

Here $\tilde{\boldsymbol{\phi}} = [\tilde{\Phi}'_{(1)}, \tilde{\Phi}'_{(2)}, \ldots, \tilde{\Phi}'_{(n)}]'$ is the $n(p+1)+n \times 1$ vector of constrained coefficients obtained from stacking the rows $\tilde{\Phi}_{(k)}$ of matrix $\tilde{\Phi}$, while $\boldsymbol{\gamma}$ is the $n(p+1)+1 \times 1$ vector of unconstrained coefficients and is defined as $\boldsymbol{\gamma} = [\Phi'_{(1)}, \Phi'_{(2)}, \ldots, \Phi'_{(n)}, \delta'_0, \delta_1]'$.

The two coefficients are linearly related through the matrix R, which is defined by

$$R = \begin{pmatrix} W & \mathbf{0} & \ldots & E_1 \\ \mathbf{0} & \ddots & \mathbf{0} & \vdots \\ \vdots & \mathbf{0} & W & E_n \end{pmatrix},$$

where W and E_k, for $k = 1, \ldots, n$ satisfy

$$W = \begin{pmatrix} I_{p\times p} \\ 0_{1\times p} \\ 0_{n\times p} \end{pmatrix} \qquad \text{and} \qquad E_k = \begin{pmatrix} 0_{p\times n} & 0_{p\times 1} \\ e'_k & 0 \\ 0_{n\times n} & e_k \end{pmatrix}.$$

The assumptions (18) on the distribution of the residuals imply that the vector of residuals $\tilde{\mathbf{u}}$ has a variance-covariance matrix that is now equal to

$$\Sigma_{\tilde{\mathbf{u}}} = \sum_q \left\{ \Sigma_{q,gr} \otimes (I_T^{gr} I_q) + \Sigma_{q,ngr} \otimes (I_T^{ngr} I_q) \right\},$$

where I_T^{gr} and I_T^{ngr} are defined as above and I_q is a $T \times T$ matrix defined as

$$I_q = \begin{pmatrix} 0_{T_1\times T_1} & \ldots & \ldots & \ldots & \ldots & \ldots & \ldots \\ \ldots & \ldots & 0_{T_{q-1}\times T_{q-1}} & 0_{T_{q-1}\times T_q} & \ldots & \ldots & \ldots \\ \ldots & \ldots & 0_{T_{q-1}\times T_q} & I_{T_q\times T_q} & 0_{T_q\times T_{q+1}} & \ldots & \ldots \\ \ldots & \ldots & \ldots & \ldots & \ldots & \ldots & \ldots \\ \ldots & \ldots & \ldots & \ldots & \ldots & \ldots & 0_{T_Q\times T_Q} \end{pmatrix}, \tag{20}$$

where $0_{T_q\times T_q}$ and $I_{T_q\times T_q}$ denote a zero matrix and an identity matrix with dimension $T_q \times T_q$, respectively.

We now derive an estimator for this model under this set of assumptions. Substituting the constraint into the main equation, the GLS estimator of $\boldsymbol{\gamma}$ minimizes the expression

$$\tilde{\mathbf{u}}'(\Sigma_{\tilde{\mathbf{u}}})^{-1}\tilde{\mathbf{u}} = \left(\mathbf{y} - (I_n \otimes \tilde{Z}')R\boldsymbol{\gamma}\right)' \Sigma_{\tilde{\mathbf{u}}}^{-1} \left(\mathbf{y} - (I_n \otimes \tilde{Z}')R\boldsymbol{\gamma}\right).$$

The estimator of $\boldsymbol{\gamma}$ is thus

$$\hat{\boldsymbol{\gamma}} = \left[R'(I_n \otimes \tilde{\mathbb{Z}})\Sigma_{\mathbf{u}}^{-1}(I_n \otimes \tilde{\mathbb{Z}}')R\right]^{-1} R'(I_n \otimes \tilde{\mathbb{Z}}')\Sigma_{\mathbf{u}}^{-1}\mathbf{y}. \tag{21}$$

Under the assumptions in (18) and the same regularity conditions as before, we can show

Claim 3 *The GLS estimator $\hat{\boldsymbol{\gamma}}$ converges in distribution according to*

$$\sqrt{T}(\hat{\boldsymbol{\gamma}} - \boldsymbol{\gamma}) \xrightarrow{d} N(\mathbf{0}, \mathcal{Q}^{-1}\mathcal{V}\mathcal{Q}^{-1}).$$

where $\mathcal{Q} \equiv \operatorname{plim}\left(T^{-1}(R'(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')R)\right)$ is well-defined, symmetric and non-singular and $\mathcal{V} \equiv R'(I_n \otimes \tilde{Z}')\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z})R$.

Proof. Rearranging (21) we obtain

$$\sqrt{T}(\hat{\boldsymbol{\gamma}} - \boldsymbol{\gamma}) = \left[\frac{R'(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')R}{T}\right]^{-1} \frac{R'(I_n \otimes \tilde{Z}')\Sigma_{\tilde{\mathbf{u}}}^{-1}\tilde{\mathbf{u}}}{\sqrt{T}}. \tag{22}$$

The expansion of the term within squared bracket yields

$$\frac{R'(I_n \otimes \tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n \otimes \tilde{Z}')R}{T} = \tag{23}$$

$$= \frac{1}{T}\begin{pmatrix} W'\tilde{Z} & 0 & \dots & 0 \\ 0 & W'\tilde{Z} & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \dots & W'\tilde{Z} \\ E_1'\tilde{Z} & E_2'\tilde{Z} & \dots & E_n'\tilde{Z} \end{pmatrix} \begin{pmatrix} \Omega_{11} & \Omega_{12} & \dots & \Omega_{1n} \\ \Omega_{21} & \Omega_{22} & \dots & \Omega_{2n} \\ \vdots & \vdots & \ddots & \vdots \\ \Omega_{n1} & \Omega_{n2} & \dots & \Omega_{nn} \end{pmatrix} \begin{pmatrix} \tilde{Z}'W & 0 & \dots & \dots & \tilde{Z}'E_1 \\ 0 & \tilde{Z}'W & \dots & \dots & \tilde{Z}'E_2 \\ \vdots & \vdots & \ddots & \dots & \vdots \\ 0 & 0 & \dots & \tilde{Z}'W & \tilde{Z}'E_n \end{pmatrix}$$

$$= \frac{1}{T}\begin{pmatrix} W'\tilde{Z}\Omega_{11}\tilde{Z}'W & W'\tilde{Z}\Omega_{12}\tilde{Z}'W & \dots & \dots & \sum_j W'\tilde{Z}\Omega_{1j}\tilde{Z}'E_j \\ W'\tilde{Z}\Omega_{21}\tilde{Z}'W & W'\tilde{Z}\Omega_{22}\tilde{Z}'W & \dots & \dots & \sum_j W'\tilde{Z}\Omega_{2j}\tilde{Z}'E_j \\ \vdots & \vdots & \ddots & \dots & \dots \\ \vdots & \vdots & \dots & W'\tilde{Z}\Omega_{nn}\tilde{Z}'W & \sum_j W'\tilde{Z}\Omega_{nj}\tilde{Z}'E_j \\ \sum_i E_i'\tilde{Z}\Omega_{i1}\tilde{Z}'W & \sum_i E_i'\tilde{Z}\Omega_{i2}\tilde{Z}'W & \dots & \sum_i E_i'\tilde{Z}\Omega_{in}\tilde{Z}'W & \sum_j \left(\sum_i E_i'\tilde{Z}\Omega_{ij}\tilde{Z}'E_j\right) \end{pmatrix}$$

where Ω_{ij} is used again to indicate the $T \times T$ block of matrix $\Sigma_{\mathbf{u}}$ with (i, j) position.

We can further rearrange this equation. To do so, we need first to realize that the matrix Ω_{ij} is now formed by different submatrices which account for the varying covariance of residuals across quarters:

$$\Omega_{ij} = \begin{pmatrix} \Omega_{ij,1} & 0 & \dots & 0 \\ 0 & \Omega_{ij,2} & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \dots & \Omega_{ij,Q} \end{pmatrix}.$$

Let $I_{gr,q}$ define the non-zero diagonal submatrix with $T_q \times T_q$ dimension of the matrix $I_T^{gr} \cdot I_q$. Matrix $I_{ngr,q}$ is defined analogously. Applying Claim 1, we write

$$\Omega_{ij,q} = \omega_{ij,gr,q} \cdot I_{gr,q} + \omega_{ij,ngr,q} \cdot I_{ngr,q},$$

where $\omega_{ij,gr,q}$ and $\omega_{ij,ngr,q}$ denote the components with (i,j) position of $\Sigma_{gr,q}^{-1}$ and $\Sigma_{ngr,q}^{-1}$, respectively.

Using the definitions of Ω_{ij}, W and each E_i, we can now show that the blocks of the expanded matrix in (23) contains the uncentered second moments of the regressors in $\tilde{z}_t$ and thus must have a well-defined probability limit. Consider a partition of the previously defined matrix $\tilde{Z}$ into a number of submatrices $\tilde{Z}_q$ which are formed by the time series of $\tilde{z}_t$ within each quarter q. Thus, $\tilde{Z} = [\tilde{Z}_1, \dots, \tilde{Z}_Q]$. We can show first that

$$\frac{1}{T} W' \tilde{Z} \Omega_{ij} \tilde{Z}' W = \frac{1}{T} Z \Omega_{ij} Z' = \sum_{q=1}^{Q} \frac{1}{T} Z_q \Omega_{ij,q} Z_q{}'$$

and, by the above decomposition of $\Omega_{ij,q}$, each element in the sum can be expressed as

$$\frac{1}{T} Z_q \Omega_{ij,q} Z_q' = n \left(\omega_{ij,gr,q} \lambda_q \sum_{t \in \mathcal{T}_q} \frac{z_t z_t'}{T_{gr,q}} + \omega_{ij,ngr,q} (1-\lambda_q) \sum_{t \in T_q \setminus \mathcal{T}_q} \frac{z_t z_t'}{T_{ngr,q}} \right),$$

where λ_q, $T_{gr,q}$ and $T_{ngr,q}$ are defined as before but now depend on a specific quarter q and $n = P(t \text{ in quarter } q)$ denotes the sample probability of a date t being in quarter q. Assuming that the limits for λ_q and n as $T \to \infty$ exist and are finite, our set of assumptions ensures that $\frac{1}{T} Z_q \Omega_{ij,q} Z_q'$ converges in probability. We can thus write

$$\frac{1}{T} W' \tilde{Z} \Omega_{ij} \tilde{Z}' W \xrightarrow{p} \mathcal{Q}_{ij},$$

where $\mathcal{Q}$ is a symmetric matrix (since $\Sigma_{\mathbf{u}}^{-1}$ is symmetric) and is assumed to be non-singular.

Using a similar rearrangement, we can also show that

$$\frac{1}{T} E_i' \tilde{Z} \Omega_{ij} \tilde{Z}' W = \sum_{q=1}^{Q} \frac{1}{T} E_i' \tilde{Z}_q \Omega_{ij,q} Z_q{}'$$

where for each q it holds

$$E_i'\tilde{Z}_q\Omega_{ij,q}Z_q{}' = n\left(\omega_{ij,gr,q}\lambda_q\sum_{t\in\mathcal{T}_q}\frac{\tilde{e}_{i,t}z_t'}{T_{gr,q}} + \omega_{ij,ngr,q}(1-\lambda_q)\sum_{t\in T_q\setminus\mathcal{T}_q}\frac{\tilde{e}_{i,t}z_t'}{T_{ngr,q}}\right).$$

Here $\tilde{e}_{i,t} \equiv E_i'\tilde{z}_t = [(e_i\cdot\mathbf{1}_{gr,t})', (e_i'\cdot(l_{gr,q}\cdot\mathbf{1}_{gr,t}))']'$ is a $(n+1)\times 1$ vector which includes the indicator $\mathbf{1}_{gr,t}$ in the i^{th} position and the interaction between the Greek indicator and the i^{th} component of vector $l_{gr,q}$ in the $(n+1)^{th}$ position. As this term has a well-defined limit under our regularity assumptions, we can conclude that the matrix $\frac{1}{T}E_i'\tilde{Z}\Omega_{ij}\tilde{Z}'W$ converges in probability and thus

$$\frac{1}{T}\sum_i E_i'\tilde{Z}\Omega_{ij}\tilde{Z}'W \xrightarrow{p} \mathcal{E}_{ij},$$

where $\mathcal{E}_{ij}$ is well-defined. Finally, we can show analougsly that

$$\sum_j\left(\sum_i E_i'\tilde{Z}\Omega_{ij}\tilde{Z}'E_j\right)\xrightarrow{p}\mathcal{H},$$

where $\mathcal{H}$ is well-defined. Combining these three results, we obtain

$$\frac{R'(I_n\otimes\tilde{Z})\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n\otimes\tilde{Z}')R}{T}\xrightarrow{p}\mathcal{Q}, \tag{24}$$

where $\mathcal{Q}$ is well-defined, symmetric and non-singular.

Given the assumption that the residual u_t is conditional independent of the regressors in $\tilde{z}_t$, we can now apply the Central Limit Theorem, as we did in the previous case, and we can show that the second term in (22) has a well-defined asymptotic distribution:

$$\frac{R'(I_n\otimes\tilde{Z}')\Sigma_{\tilde{\mathbf{u}}}^{-1}\tilde{\mathbf{u}}}{\sqrt{T}}\xrightarrow{d}N(\mathbf{0},\mathcal{V}), \tag{25}$$

where $\mathcal{V}\equiv R'(I_n\otimes\tilde{Z}')\Sigma_{\tilde{\mathbf{u}}}^{-1}(I_n\otimes\tilde{Z})R$. Combining (24) and (25) we obtain

$$\sqrt{T}(\hat{\boldsymbol{\gamma}}-\boldsymbol{\gamma})\xrightarrow{d}N(\mathbf{0},\mathcal{Q}^{-1}\mathcal{V}\mathcal{Q}^{-1}).$$

■

Table* 1*: Pairwise Correlations of Changes in Sovereign CDS for* 11 *European Countries

***Log changes by business day, January* 1 2009 *to December* 31 2010**

	Austria	Belgium	France	Germany	Greece	Italy	Netherlands	Portugal	Spain	Sweden
Belgium	0.7678									
France	0.7470	0.7832								
Germany	0.7322	0.7395	0.8166							
Greece	0.6545	0.6809	0.6813	0.6200						
Italy	0.7425	0.7693	0.7516	0.7120	0.7426					
Netherlands	0.7595	0.8002	0.7475	0.7361	0.6285	0.7370				
Portugal	0.7108	0.7396	0.7425	0.6830	0.8005	0.8420	0.7059			
Spain	0.7253	0.7707	0.7363	0.7209	0.7829	0.8952	0.7190	0.8845		
Sweden	0.7700	0.7152	0.7019	0.7073	0.6011	0.6728	0.7322	0.6469	0.6721	
UK	0.7444	0.7584	0.7178	0.6731	0.6118	0.7434	0.7404	0.7244	0.7147	0.6962

Source: Datastream, own calculations.

Table* 2a*: Summary Statistics - Exposure to Greek Debt, Time-Average
in % of GDP, Average 2009/10

	Total Debt	**Public Debt**	**Bank Debt**
mean	0.74%	0.48%	0.09%
std. dev.	0.46%	0.31%	0.07%
min	0.08%	0.04%	0.00%
max	1.22%	0.77%	0.21%
No. Countries	11	11	11

Source: BIS Consolidated Banking Statistics. Countries are: Austria, Belgium, France, Germany, Italy, the Netherlands, Portugal, Spain, Sweden and the UK.

Table* 2b*: Summary Statistics - Exposure to Greek Debt, Evolution
***Quarterly data, normalized* 2009*Q*4 = 100**

			A.Total Debt		
	2009Q4	2010Q1	2010Q2	2010Q3	2010Q4
mean	100.00%	107.10%	81.50%	78.80%	67.50%
std. dev.	0.00%	18.90%	17.60%	18.20%	24.80%
min	100.00%	82.20%	48.80%	42.60%	22.80%
max	100.00%	152.90%	110.70%	101.20%	99.80%
			B. Public Debt		
	2009Q4	2010Q1	2010Q2	2010Q3	2010Q4
mean	100%	116.3%	88.0%	85.3%	73.0%
std. dev.	0%	30.8%	30.9%	31.7%	34.9%
min	100%	85.5%	23.9%	17.6%	17.7%
max	100%	178.1%	123.3%	121.7%	113.1%
			C. Bank Debt		
	2009Q4	2010Q1	2010Q2	2010Q3	2010Q4
mean	100%	92.5%	77.9%	59.1%	32.0%
std. dev.	0%	58.3%	77.4%	74.2%	30.4%
min	100%	9.4%	19.1%	7.9%	2.3%
max	100%	229.9%	277.8%	265.8%	104.2%
No. Countries	10	10	10	10	10

Source: BIS Consolidated Banking Statistics. Countries are: Austria, Belgium, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain, Sweden and the UK.

Table 3a: Summary Statistics - Greek CDS spreads
Log changes by business day, January 1 2009 to December 31 2010

	All days, excluding Greek events	Days of positive Greek shocks	Days of negative Greek shocks
mean	0.27%	7.37%	-7.51%
std. dev.	4.98%	8.61%	5.15%
min	-47.10%	-0.67%	-15.70%
max	19.30%	27.00%	-2.33%

Source: Datastream, own calculations

Table 3b: Summary Statistics - CDS spreads in 10 European countries during Greek events
Log-changes by business day, by type and level of exposure to Greek debt.

A. Total Debt

	Days of positive Greek shocks		Days of negative Greek shocks	
	Exposure above country-median	Exposure below country-median	Exposure above country-median	Exposure below country-median
mean	3.03%	1.85%	-2.99%	-0.49%
std. dev.	4.42%	4.73%	3.36%	3.04%
min	-3.51%	-4.75%	-7.76%	-6.85%
max	17.70%	13.30%	1.94%	2.44%

B. Public Debt

	Days of positive Greek shocks		Days of negative Greek shocks	
	Exposure above country-median	Exposure below country-median	Exposure above country-median	Exposure below country-median
mean	3.05%	2.76%	-3.10%	-0.48%
std. dev.	4.88%	3.93%	3.28%	2.76%
min	-4.75%	-2.69%	-7.76%	-6.85%
max	17.70%	11.40%	1.94%	2.44%

C. Bank Debt

	Days of positive Greek shocks		Days of negative Greek shocks	
	Exposure above country-median	Exposure below country-median	Exposure above country-median	Exposure below country-median
mean	3.67%	1.58%	-2.18%	-0.97%
std. dev.	4.72%	3.94%	3.62%	3.05%
min	-3.51%	-4.75%	-7.76%	-7.48%
max	17.70%	11.40%	5.84%	2.26%

Source: Datastream, BIS Consolidated Banking Statistics. Countries are: Austria, Belgium, France, Germany, Italy, the Netherlands, Portugal, Spain, Sweden and the UK.

Table 4a: ***Baseline GLS Estimation Results - logged CDS***

	AT	BE	DE	ES	FR	GB	IT	NL	PT	SE	GR
GR	0.0160	0.0183	0.0036	0.0221	0.0233	0.0184	0.0277	0.0074	0.0243	0.0183	0.0631
	0.0096	0.0113	0.0114	0.0110	0.0071	0.0086	0.0101	0.0091	0.0119	0.0079	0.0138
Rsq	0.53	0.55	0.56	0.57	0.61	0.54	0.59	0.55	0.60	0.48	0.54
Obs.	492	492	492	492	492	492	492	492	492	492	492
RResp	0.2536	0.2900	0.0571	0.3502	0.3693	0.2916	0.4390	0.1173	0.3851	0.2900	
					Bootstrap Results						
mean (bootstr)	0.242	0.2788	0.0318	0.3615	0.3675	0.3036	0.4438	0.1025	0.3809	0.2915	
.005 per-tile	-0.1853	-0.2173	-0.6954	-0.1641	0.1503	-0.0683	0.0209	-0.4056	-0.1111	-0.0027	
.025 per-tile	-0.0559	-0.0746	-0.4517	0.013	0.1986	-0.0008	0.1157	-0.2586	0.0183	0.0748	
.975 per-tile	0.472	0.5606	0.3814	0.7499	0.5584	0.6331	0.7693	0.3613	0.7254	0.5038	
.995 per-tile	0.5174	0.6344	0.4837	0.9215	0.6275	0.8272	0.8988	0.4203	0.8569	0.5768	

Note: Dependent Variable vector of CDS spreads, in logs. Countries are: Austria, Belgium, Germany, Spain, France, UK, Italy, the Netherlands, Portugal, Sweden and Greece. The underlying model is as specified in equation (6). Exogenous variables are CDS for US and Japan (log), VIX of the US (log), federal fund rate, S&P market returns, and weekday dummies. Eight lags of the dependent variables are included, none of the exogenous variables. All CDS spreads and the VIX are in first differences to remove a stochastic trend.

Table 4b: ***Baseline GLS Estimation Results - CDS in basis points***

	AT	BE	DE	ES	FR	GB	IT	NL	PT	SE	GR
GR	1.105	0.975	-0.012	3.078	1.075	1.434	3.557	0.403	5.015	0.872	30.155
	0.816	0.929	0.478	1.737	0.411	0.702	1.313	0.413	3.414	0.408	8.525
Rsq	0.54	0.55	0.61	0.58	0.62	0.56	0.61	0.54	0.59	0.51	0.53
Obs.	492	492	492	492	492	492	492	492	492	492	492
RResp	0.0366	0.0323	-0.0004	0.1021	0.0356	0.0475	0.1180	0.0134	0.1663	0.0289	
					Bootstrap Results						
mean (bootstr)	0.034	0.0295	-0.0027	0.1027	0.0357	0.0515	0.1222	0.0122	0.1547	0.03	
.005 per-tile	-0.0808	-0.1357	-0.1045	-0.091	0.004	-0.0208	0.0178	-0.0415	-0.2304	-0.0064	
.025 per-tile	-0.0269	-0.0446	-0.054	-0.0097	0.0157	0.0037	0.0419	-0.0196	-0.0895	0.0029	
.975 per-tile	0.0793	0.0869	0.0322	0.2148	0.0554	0.1205	0.2116	0.0387	0.3523	0.0612	
.995 per-tile	0.0975	0.1075	0.0444	0.253	0.0656	0.1564	0.2482	0.0506	0.3821	0.0857	

Note: Dependent Variable vector of CDS spreads, in levels (bp). Countries are: Austria, Belgium, Germany, Spain, France, UK, Ireland, Italy, the Netherlands, Portugal, Sweden and Greece. The underlying model is as specified in equation (6). Exogenous variables are CDS for US and Japan, in levels (bp), VIX of the US, federal fund rate, S&P market returns, and weekday dummies. Eight lags of the dependent variables are included, non of the exogenous variables. All CDS spreads and the VIX are in first differences to remove a stochastic trend.

Table 5: ***Financial Linkages - Logged Exposure***

	I	II	III
GR	0.0679	0.0675	0.0674
	0.0134	0.0134	0.0135
GR*Total	0.0376		
	0.0081		
GR*Publ		0.0266	
		0.0049	
GR*Bank			0.0099
			0.0020
***d* (Total)**	0.5538		
***d* (Public)**		0.3941	
***d* (Bank)**			0.1469
		Bootstrap Results	
mean (bootstr)	0.5415	0.4022	0.1308
.005 per-tile	0.2775	0.2179	0.0568
.025 per-tile	0.3142	0.2501	0.0713
.975 per-tile	0.8514	0.6189	0.2166
.995 per-tile	1.0735	0.7401	0.2558

Note: Dependent variable vector of CDS, in logs. Countries are: Austria, Belgium, Germany, Spain, France, UK, Italy, the Netherlands, Portugal, Sweden and Greece. The underlying model is as specified in equation (11). Exogenous variables are CDS for US and Japan (log), VIX of the US (log), federal fund rate, S&P market returns and weekday dummies. Data on exposure to Greek debt are logged. Eight lags of the dependent variables are included, none of the exogenous variables. All CDS spreads and the VIX are in first differences to remove a stochastic trend.

Table 6*: *Financial Linkages - Percentage Exposure

	I	II	III
GR	35.1402	35.2130	34.8378
	8.4329	8.4326	8.4332
GR*Total	397.0294		
	72.5000		
GR*Publ		429.4027	
		73.1011	
GR*Bank			1395.4000
			238.8000
***d* (Total)**	11.2984		
***d* (Public)**		12.1944	
***d* (Bank)**			40.0542
		Bootstrap Results	
mean (bootstr)	10.7031	11.7488	38.0809
.005 per-tile	5.6408	6.5751	18.7495
.025 per-tile	6.8993	7.5114	21.8475
.975 per-tile	16.0139	17.8672	65.8435
.995 per-tile	18.2967	20.6559	75.9644

Note: Dependent variable vector of CDS, in levels (bp). The model is specified in equation (11). Countries are: Austria, Belgium, Germany, Spain, France, UK, Italy, the Netherlands, Portugal, Sweden and Greece. Exogenous variables are CDS for US and Japan (in levels), VIX of the US (in levels), federal fund rate, S&P market returns and weekday dummies. Exposure to Greek debt is expressed as ratios of creditor's GDP. See also note of Table 5.

Table 7: Financial Linkages - Excluding UK and Sweden

	I	II	III
GR	0.0674	0.0669	0.0668
	0.0148	0.0148	0.0148
GR*Total	0.0374		
	0.0108		
GR*Publ		0.0207	
		0.0058	
GR*Bank			0.0096
			0.0021
***d* (Total)**	0.5549		
***d* (Public)**		0.3094	
***d* (Bank)**			0.1437
		Bootstrap Results	
mean (bootstr)	0.5514	0.3020	0.1154
.005 per-tile	0.2003	0.1296	0.0335
.025 per-tile	0.2761	0.1535	0.0497
.975 per-tile	0.9039	0.5386	0.208
.995 per-tile	1.0166	0.635	0.2516

Note: Dependent variable vector of CDS, in logs. The model is specified in equation (11). Countries are: Austria, Belgium, Germany, Spain, France, Italy, the Netherlands, Portugal and Greece. Data on exposure to Greek debt are logged. See also Note of Table 5.

Table* 8: *Financial Linkages - Lagged Exposure

	I	II	III
GR	0.0677	0.0676	0.0679
	0.0134	0.0134	0.0134
GR*Total	0.0355		
	0.0090		
GR*Publ		0.0167	
		0.0054	
GR*Bank			0.0075
			0.0024
***d* (Total)**	0.5244		
***d* (Public)**		0.2470	
***d* (Bank)**			0.1105
		Bootstrap Results	
mean (bootstr)	0.4625	0.2094	0.1004
.005 per-tile	0.1582	0.042	0.0284
.025 per-tile	0.2319	0.0778	0.0449
.975 per-tile	0.7513	0.3839	0.1668
.995 per-tile	0.8862	0.4588	0.2023

Note: Dependent variable vector of CDS, in logs. The model is specified in equation (11). Countries are: Austria, Belgium, Germany, Spain, France, UK, Italy, the Netherlands, Portugal, Sweden and Greece. Data on exposure to Greek debt are lagged and logged. See also Note of Table 5.

Table 9: ***Financial Linkages - Indirect Linkages***

	I	II	III
GR	0.0676	0.0674	0.0672
	0.0134	0.0134	0.0134
GR*Total	0.0486		
	0.0129		
GR*Publ		0.0331	
		0.0059	
GR*Bank			0.0173
			0.0034
***d* (Total)**	0.7189		
***d* (Public)**		0.4911	
***d* (Bank)**			0.2574
		Bootstrap Results	
mean (bootstr)	0.7389	0.5098	0.2420
.005 per-tile	0.3669	0.2627	0.1143
.025 per-tile	0.4129	0.3161	0.138
.975 per-tile	1.2093	0.7928	0.381
.995 per-tile	1.417	0.9611	0.4486

Note: Dependent variable vector of CDS, in logs. The model is specified in equation (11). Countries are: Austria, Belgium, Germany, Spain, France, UK, Italy, the Netherlands, Portugal, Sweden and Greece. Indirect exposures to Greek debt are obtained by inverting the matrix of cross-border bilateral linkages and logged. See also Note of Table 5.

Table* 10: *Financial Linkages - Total Exposure to Foreign Debt

	I	II	III
GR	0.0683	0.0677	0.0683
	0.0134	0.0134	0.0134
GR*Total	0.0062		
	0.0282		
GR*Publ		-0.0261	
		0.0167	
GR*Bank			0.0118
			0.0174
***d* (Total)**	0.0908		
***d* (Public)**		-0.3855	
***d* (Bank)**			0.1728
		Bootstrap Results	
mean (bootstr)	-0.0167	-0.4584	0.0953
.005 per-tile	-0.9543	-1.2832	-0.4478
.025 per-tile	-0.7461	-1.0214	-0.3191
.975 per-tile	0.7187	-0.0425	0.5249
.995 per-tile	0.9338	0.0804	0.6425

Note: Dependent variable vector of CDS, in logs. The model is specified in equation (11). Countries are: Austria, Belgium, Germany, Spain, France, UK, Italy, the Netherlands, Portugal, Sweden and Greece. Total exposure to foreign debt (in logs) is the sum of a country's exposure to any other country in the BIS dataset. See also Note of Table 5.

Table **A1***: Financially Relevant Information Shocks of Greek Origin*

Date	Description	Sign
10/9/2009	Greece's central bank governor said the country's budget deficit this year may exceed 12 percent of economic output or four times the EU ceiling. George Provopoulos said current data for the first three quarters of 2009 set budget overspending at around 10 percent of gross domestic product. (Source: Lexis Nexis)	+
12/8/2009	Fitch Ratings cuts Greek debt to BBB+, the first time in 10 years it has been rated below investment grade. (Source: Reuters)	+
3/3/2010	The Greek government announces a new austerity plan that will yield €4.8 billion in savings. The government decides to move ahead with steep cuts in civil-service salaries and entitlements, and to to raise Greece's sales tax by two percentage points. (Source: Wall Street Journal)	-
3/11/2010	An estimated 50,000 people take to the streets to protest the government's austerity plans. Flights were grounded and trains suspended amid a nationwide general strike. (Source: Wall Street Journal)	+
3/25/2010	Mr. Trichet says the ECB will continue to accept bonds with ratings as low as triple-B-minus as collateral next year, although it won't lend as much against lower-rated paper as against the ultra-safe triple-A-rated bonds like Germany's. (Source: Wall Street Journal)	-
4/12/2010	Eurozone members commit to provide up to 30bn euros in loans to Greece over the next year. The funds would be supplemented by money from the IMF that could yield an additional 15bn euros. The rates charged to Athens would be about 5 percent for a three-year fixed loan. (Source: Financial Times)	-
4/22/2010	Eurostat said Greece's budget deficit reached 13.6 percent of GDP last year, up from previous Eurostat estimates of 12.9 percent. The news came as thousands of Greek civil servants staged a 24-hour strike and a march through central Athens on Thursday in protest gainst planned reforms of the state-funded pension system due to reach parliament next month. (Source: Financial Times)	+
5/2/2010	Prime Minister George Papandreou says Greece has reached a deal with the EU and IMF opening the door to a bailout in return for extra savings of 30 billion euros over three years. Athens will get loans worth 110 billion euros in instalments conditional on reforms over three years in the first rescue of a member of the 16-nation euro zone. (Source: Reuters)	-
5/3/2010	The ECB suspends the minimum credit rating required for Greek government-backed assets used in its liquidity-rpoviding operations, removing the risk of Greek government bonds being excluded if ratings agencies turn against the rescue programme. The euro fell 0.9 percent against the US dollar, although bond markets across Europe were stable with yields on two-year Greek debt dropping below 10 percent for the first time in more than a week. Speculators are increasingly betting that the eurozone crisis could escalated as short positions against the euro rose to a fresh record. (Source: Financial Times)	-
5/20/2010	Thousands of Greek workers protest in Athens against cuts in wages and pensions. The 24-hour general strike called by ADEBY and GSEE, the country's two largest unions, has shut down state schools, public transport and government offices. Parliament is set to approve an overhaul of the state pension system this month under the terms of the bail-out. The legislation includes proposal to raise the retirement age to 65 for both men and women from 2013 and reduce pension payments. (Source: Financial Times)	+
6/15/2010	The ECB says it will apply a 5 percent extra charge to Greek government bonds used as collateral in lending operations following a downgrade. (Source: Reuters)	+
8/12/2010	New statistics suggest Greek sank deeper into recession in the second quarter, with Elstat, the statistic service, estimating the economy had shrunk 3.5. percent in the three months to the end of June from the same period last year. There was also sharp rise in the year-on-year jobless rate from 8.5. percent to 12.5 percent in May a record increase reflecting worsening conditions in the real economy, according to analysts. (Source: Financial Times)	+
9/8/2010	Greek second-quarter gross domestic product is revised down to show a 1.8% fall from the first three months of the year compared with the initial estimate of a 1.5% decline. (Source: Wall Street Journal)	+
10/4/2010	Greece unveils an ambitious draft budget aiming to slash its budget deficit to 7 percent of GDP next year, deeper than the 7.6 percent goal agreed with the Ecm, the ECB and the IMF and aimed at securing a return to the bond market at some point next year. The 2011 draft budget sees the general government budget deficit drop to 16.28 bn in 2011 from a forecast 18.5 bn this year on higher tax revenues and spending cuts with real GDP projected to contract by 2.6 percent. Additionally, Greek 10-year bond yields fell 15 basis points to 9.975 percent following comments on Sunday by Wen Jiabao, China premier, on Bejing interest in buying Greek bonds. (Source: Financial Times)	-